University of Oslo

Doctoral Thesis

Housing markets and financial stability

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Oslo, November 2013,

André Kallåk Anundsen

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To Dunja,

Thanks for your patience and support

Chapter 1

Introduction

What goes up must come down? Starting in the late 1990s, there was an unprecedented international housing price boom accompanying the favorable economic situation in most industrialized countries, with Germany, Switzerland and Japan being the major exceptions.[1] It is clear from Figure 1.1 – where I have plotted the real housing price development in 19 industrialized countries since 1990 – that the boom was succeeded by a significant bust in many countries, with real housing prices falling by more than 30 percent in several cases. The consequences for the real economy following the bust in housing prices have been severe and it was one of the factors contributing to the deepest downturn in the world economy since the Great Depression. In countries such as Ireland and Spain, the unemployment rates in the construction sector rose to record levels as investments plummeted. The collapse culminated with the meltdown of the US housing market and financial system in 2007/2008 – the epicenter for the ensuing global financial crisis that still puts a strain on global economic recovery. In the US, the collapse triggered a massive deleveraging process and the savings rate tripled during the Great Recession (see Glick and Lansing (2009)). The real economic consequences have been severe, and Lansing (2011) has estimated the *per capita* foregone consumption during the period from late 2007 to May 2011 to be $7,300.

Against this background, it should be clear that a good understanding of the interaction between the real economy and the financial markets is key to monitor the stability of the real economy and the financial system, and it is important for the conduct of both monetary and regulatory policies, see e.g. the discussion in Muellbauer (2010). In that respect, the development in the housing market is of particular relevance, since a housing

[1] Gros (2007) argues that the moderate price development in Germany is largely a result of the excess supply resulting from a high building activity in the years after the reunification of East and West Germany, while Japan has gone through its lost decade (Kim and Renaud, 2009) and Switzerland has an unusual market structure with very low home ownership rates (Bourassa and Hoesli, 2010).

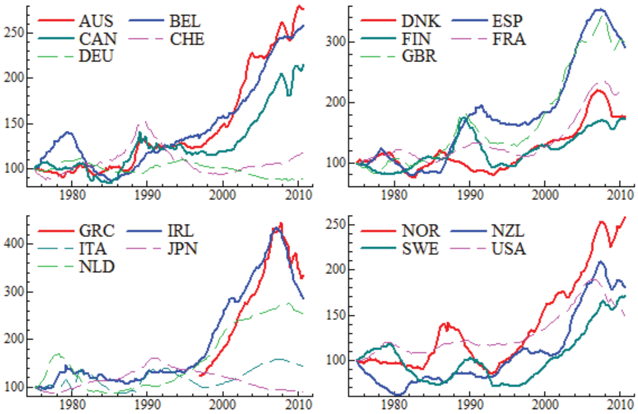


FIGURE 1.1: Real housing price development in 19 OECD countries. *Notes:* The following abbreviations apply. AUS = Austria, BEL = Belgium, CAN = Canada, CHE = Switzerland, DEU = Germany, DNK = Denmark, ESP = Spain, FIN = Finland, FRA = France, GBR = Great Britain, GRC = Greece, IRL = Ireland, ITA = Italy, JPN = Japan, NLD = Netherlands, NOR = Norway, NZL = New Zealand, SWE = Sweden and USA = United States of America. (*Source:* OECD)

purchase is the single biggest investment made by most households during the course of a life, and it constitutes the major slice of household wealth.

The housing market may have important feedback effects to the macro economy and the financial system, and the development in the housing market may affect the real economy through both consumption wealth effects, see e.g. Brodin and Nymoen (1992) and Jansen (2013) for evidence of wealth effects on consumption in Norway and Aron et al. (2012) for evidence in the UK and the US[2], and by stimulating housing investments through a Tobin-Q effect (Tobin, 1969). In addition, most housing loans are collateralized by the property itself, which may give rise to spill-over effects between housing prices and household borrowing. There are thus several channels in which both fundamental and non-fundamental movements in housing prices may jeopardize the soundness of the real economy and the entire financial system (see also the discussion in Goodhart and Hofmann, 2007). With this in mind, it is interesting to note that Borio and Lowe (2002) find that there are several cases where increasing housing and stock prices together with a credit expansion have signaled an increasing financial and real economic instability, and – as pointed out by Koetter and Poghosyan (2010) and Goodhart and Hofmann (2007) – there are numerous episodes where falling housing prices have preceded financial and banking crises in a historical context. This is one of the reasons why

[2]It should be noted that the same authors find a negative housing wealth effect for Japan. Aron et al. (2012) attribute this to the absence of credit market liberalization.

policymakers keep a close eye at the price development in the housing market when assessing the vulnerability of the financial system. Furthermore, Leamer (2007) find that 8 out of the 10 post World War II recessions in the US have been preceded by a decline in housing construction and durable consumption, suggesting that housing starts is a good leading indicator for the future economic development – a claim that parallels the findings of Davis and Heathcote (2005).

Both the Norwegian banking crisis in the early 1990s and the recent financial crisis are examples of how a growing instability in the housing and credit markets have threatened the stability of the financial system, with huge consequences for the real economy. Figure 1.2 plots the run-up in household leverage (loan-to-income) during the pre financial crisis period between 1997–2007 against the percentage change in private consumption over the 2008–2009 period for 16 industrialized countries.[3] Though no causal conclusions should be derived from this simple scatter plot, it is clear that the countries that had the highest leveraged households prior to the crisis are the same countries that experienced the greatest decline in private consumption during the crisis.

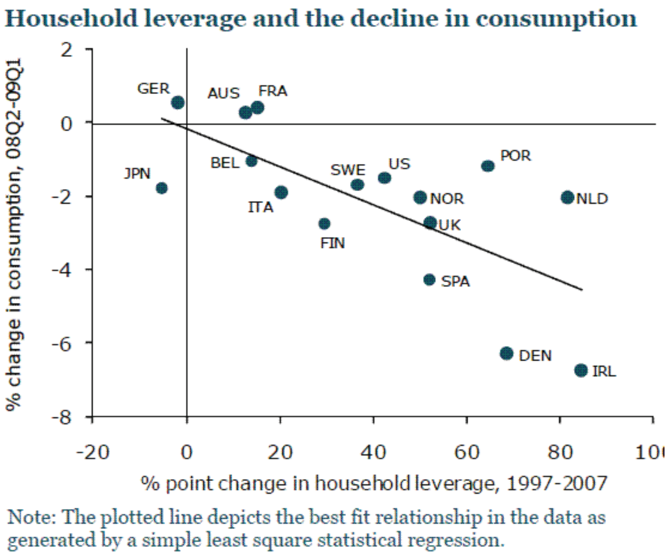


FIGURE 1.2: Household leverage and the decline in consumption. (Source: Glick and Lansing (2010))

Recent experiences suggest that there is a great need to enhance research in the area of housing economics, and in particular on the interaction between the housing market, credit markets and the real economy. This thesis contributes in that respect by exploring a range of issues in this area, such as the interaction between housing markets and credit

[3]This figure is taken from Glick and Lansing (2010). Thanks to Kevin Lansing for sharing the figure.

markets, by suggesting a methodological framework in which pending imbalances in the housing market may be detected in real time, and by exploring what factors contribute to explain regional differences in housing price dynamics and long-run housing price determination.

The rest of this introduction proceeds as follows. First, I discuss the theory of housing demand and supply, and provide an overview of a selection of the many important research areas within the field of housing economics, while also connecting the contributions of this thesis to that literature. The final part of the introduction summarizes the four papers that comprise this thesis.

1.1 Theory of the housing market

1.1.1 Housing demand and the relationship between housing prices and rents

A central building block underpinning large parts of the econometric modeling carried out in the different chapters of this thesis is the theory of housing demand. The most commonly used framework in empirical housing studies is the life-cycle model of housing (see e.g. the seminal contribution of Dougherty and Van Order (1982)), which is well founded in microeconomic theory.

Consider a representative consumer that maximizes his lifetime utility with respect to housing consumption, H, and consumption of "other goods", C. The discount factor is given by β, and utility is maximized subject to a budget constraint and two technical constraints describing the law of motion of housing capital and net non-housing wealth, W, respectively:

$$max \int_0^\infty e^{-\beta t} u(C_t, H_t) dt$$

subject to:

$$PH_t I_t + S_t + C_t = (1 - \theta_t) Y_t + (1 - \theta_t) i_t W_t$$
$$\dot{H}_t = I_t - \delta H_t$$
$$\dot{W}_t = S_t - \pi_t W_t$$

where PH denotes real housing prices, S is net real savings (savings net of new loans), I is investments in new housing capital, θ is the marginal tax rate, Y is real household income and i and π denote the nominal interest rate and the CPI inflation, respectively. Thus, the budget constraint states that the sum of expenditures on housing and other consumer goods plus savings is equal to after tax income plus the interest earned on net non-housing wealth. The law of motion of capital equation says that new housing is given by new investments less the depreciation of the existing stock. The law of motion of wealth says that changes in wealth are given by net-savings less depreciation of the real value of existing wealth due to changes in the overall inflation rate.

Formulating the Hamiltonian and solving the constrained maximization problem results in the following equilibrium condition (see Appendix 1.A for details):

$$\frac{U_H}{U_C} = PH_t \left[(1-\theta_t)i_t - \pi_t + \delta - \frac{\dot{PH_t}}{PH_t} \right] \tag{1.1}$$

which simply states that the marginal rate of substitution between housing and the composite consumption good is equal to what it costs to own one more unit of a property. Since the housing market also contains a rental sector, market efficiency requires the following condition to be satisfied in equilibrium:

$$Q_t = PH_t \left[(1-\theta_t)i_t - \pi_t + \delta - \frac{\dot{PH_t}}{PH_t} \right]$$

where Q_t is the real imputed rent on housing services. Hence, the price-to-rent ratio is proportional to the inverse of the user cost:

$$\frac{PH_t}{Q_t} = \frac{1}{UC_t} \tag{1.2}$$

where the user cost is defined as $UC_t = (1-\theta_t)i_t - \pi_t + \delta - \frac{\dot{PH_t}}{PH_t}$. The real imputed rent is unobservable, but two approximations are common: to proxy the imputed rent by an observable rent R_t, or to assume that it is proportional to income and the stock of housing. Relying on the first approximation, the expression in (1.2) would read:

$$\frac{PH_t}{R_t} = \frac{1}{UC_t} \tag{1.3}$$

while if we instead assume that the imputed rent is determined by the following expression:

$$R_t = Y_t^{\beta_y} H_t^{\beta_h},\ \beta_y > 0 \ and \ \beta_h < 0$$

(1.2) would read:

$$\frac{PH_t}{Y_t^{\beta_y} H_t^{\beta_h}} = \frac{1}{(1-\theta_t)i_t - \pi_t + \delta - \frac{\dot{PH_t}}{PH_t}} \tag{1.4}$$

The expressions represented by (1.3) and (1.4) are commonly used as a starting point for building econometric models of housing price formation, and they will also be central to the econometric modeling carried out in this thesis. While the first has been used extensively in the US literature, it is less common in Europe, since the rental market is relatively small in countries such as e.g. the UK and Norway, and since the rental market is heavily regulated in many European countries (Muellbauer, 2012). The expression in (1.4) is similar to an inverted demand equation, and we now see how it can be derived from a life-cycle model of housing.

A natural starting point for an econometric analysis of housing price determination is therefore to consider these expressions on a semi-logarithmic form,[4] which gives:

$$ph_t = \beta_r r_t + \beta_{UC} UC_t \tag{1.5}$$

$$ph_t = \beta_y y_t + \beta_h h_t + \beta_{UC} UC_t \tag{1.6}$$

where we would expect that $\beta_r, \beta_y > 0$ and $\beta_h, \beta_{UC} < 0$. Either or both of these equations form the basis for a series of papers that investigate housing price determination, see e.g. Buckley and Ermisch (1983); Hendry (1984); Meen (1990); Holly and Jones (1997); Muellbauer and Murphy (1997); Meen and Andrew (1998); Meen (2001); Duca et al. (2011a,b) to mention a few of the many empirical studies that are grounded in the life-cycle model of housing.

Extensions of the simple version of the life-cycle model of housing presented in this section include an explicit role of credit constraints, as in Dougherty and Van Order (1982), Meen (1990, 2001) and Meen and Andrew (1998). In that case, the expression in (1.1) would be augmented with an additional term reflecting the shadow price on a mortgage credit constraint. This would of course also entail that the expressions in (1.5) and (1.6) would be augmented with an additional term for credit constraints. As will become evident throughout this thesis, the role of such credit constraints are indeed very important for housing price formation. That said, the credit constraint variable is hard to observe, and some house buyers will always be credit constrained, while others will never be. Further, the composition of which borrowers are credit constrained and which are not – and hence the average value of this variable – may well change over time, which has important consequences for the determination of housing prices.

1.1.2 Housing supply

Why don't we know more about housing supply? This question was raised by DiPasquale in a paper from 1999 (DiPasquale, 1999). Now, almost fifteen years later, we must ask

[4] A semi-logarithmic representation is commonly used, since the user cost may take negative values.

the same question. As emphasized by Muellbauer and Murphy (2008), the empirical evidence on residential investments is very diverge with conflicting results. That said, a common starting point is to assume that investments are determined in accordance with a Tobin Q (Tobin, 1969) theory of housing, which simply states that housing investments are proportional to the ratio of the market price of existing houses to its replacement cost – which can be considered as the sum of construction costs and land costs[5], i.e.:

$$I_t = \left(\frac{PH_t}{PJ_t}\right)^{\eta} \tag{1.7}$$

where η is the elasticity of supply and PJ is a measure of the replacement cost. The expression in (1.7) is interesting in several respects. First, it provides a background to understand why some housing studies substitute the housing stock measure in (1.4) by some measure of construction costs. To see this, remember that the law of motion of capital is given by:

$$\dot{H}_t = I_t - \delta H_t$$

which means that in a static long-run equilibrium, we have:

$$H_t = \frac{1}{\delta} I_t$$

This implies that the long-run supply curve takes the following form:

$$H_t = \frac{1}{\delta}\left(\frac{PH_t}{PJ_t}\right)^{\eta} \tag{1.8}$$

Substituting (1.8) for H_t in (1.4) and considering the semi-logarithmic representation gives:

$$ph_t = \gamma_0 + \gamma_y y_t + \gamma_{pj} pj_t + \gamma_{UC} UC_t \tag{1.9}$$

Thus, we have a rationale for this alternative operationalization, which may be interpreted as a reduced form housing price equation. It is an equation of the form (1.9) that forms the basis for the analysis in e.g. Abraham and Hendershott (1996) and Berlinghieri (2010) who study US housing price determination over the period 1977–1992 and 1977–2005, respectively. Whether the researcher chooses to consider an inverted demand equation of the type (1.6) or a reduced form housing price equation of the form (1.9), theory clearly demonstrates the importance of taking into account some supply side measure when modeling housing prices.

[5] Mayer and Somerville (2000) follow another approach and argue that investments are determined by the changes in housing prices and construction costs, rather than the levels of these variables.

In addition to providing a rationale for why some studies consider the construction costs as opposed to the supply of dwellings in models of housing price determination, (1.7) clearly demonstrates the importance of the supply elasticity for the dynamics of the housing market. Consider an increase in income. From (1.4), we know that this will put an upward pressure on housing prices. However, part of the initial increase in housing prices will be dampened in the long-run, since higher housing prices leads to increased supply because of a higher investment activity (confer (1.7) and (1.8)). Thus, the higher is the supply elasticity, the lower will the total increase in housing prices following a given increase in income be.

1.2 Housing prices and credit markets

One motivation to study the housing market may be found in the theoretical literature on financial accelerators, see e.g. Bernanke and Gertler (1989) and Kiyotaki and Moore (1997).[6] The idea behind the financial accelerator in a housing market context is that imperfections in the credit markets necessitates the need for collateral when a housing loan is granted. Consequentially, imbalances in the financial markets may generate and amplify imbalances in the real economy, and *vice versa*. An increase in housing prices have both direct and indirect effects on credit, as illustrated in Figure 1.3.

As a direct effect, higher property prices increase the amount of credit needed to finance a given housing purchase. Many indirect effects are present as well, the most important probably being that an increase in housing prices leads to a higher value of borrowers' collateral, increasing their borrowing capacity. In addition to this, higher property valuations increase the value of banks' assets, thereby improving their capital position. Finally, expected life-time wealth may increase as a result of higher housing prices, leading to a greater demand for credit in order to smooth consumption over the life-cycle. On the other hand, more credit in circulation implies that the demand for housing services will, *ceteris paribus*, increase. For this reason, we see how shocks in one of these markets might transmit to the other, and thereby explaining the simultaneous occurrence of boom and bust cycles in the housing and credit markets.

[6]For literature on the financial accelerator in the context of DSGE models, confer for example Aoki et al. (2004), Iacoviello (2005) and Iacoviello and Neri (2010). Since this thesis is confined to econometric assessments of the housing market, a detailed description of this branch of the literature is beyond the scope of this introduction.

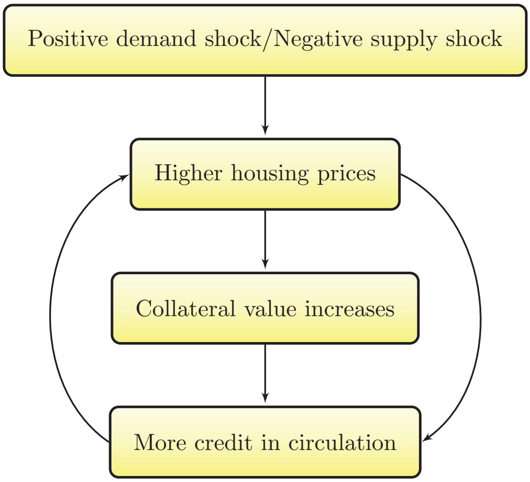


Figure 1.3: Two-way-interaction between housing prices and credit

Looking at Figure 1.4, it is clear that the countries with the most leveraged households during the 1997–2007 period also were the countries that experienced the greatest build-up of housing prices over this period, which is in accordance with a financial accelerator view.[7]

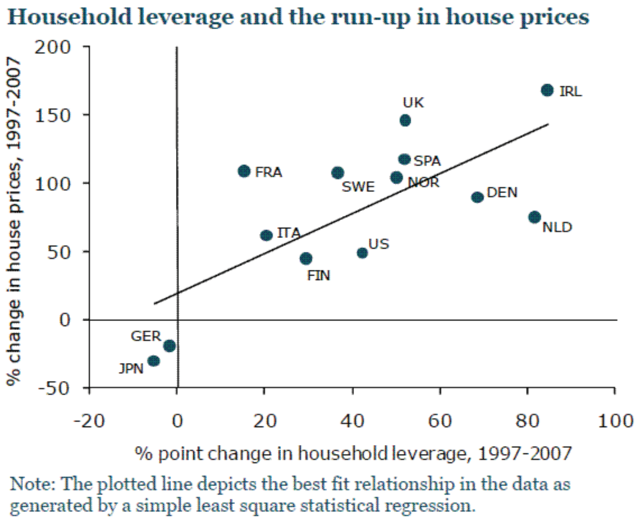


Figure 1.4: Household leverage and the run-up in house prices (Source: Glick and Lansing (2010))

[7]This figure is taken from Glick and Lansing (2010). Thanks to Kevin Lansing for sharing the figure.

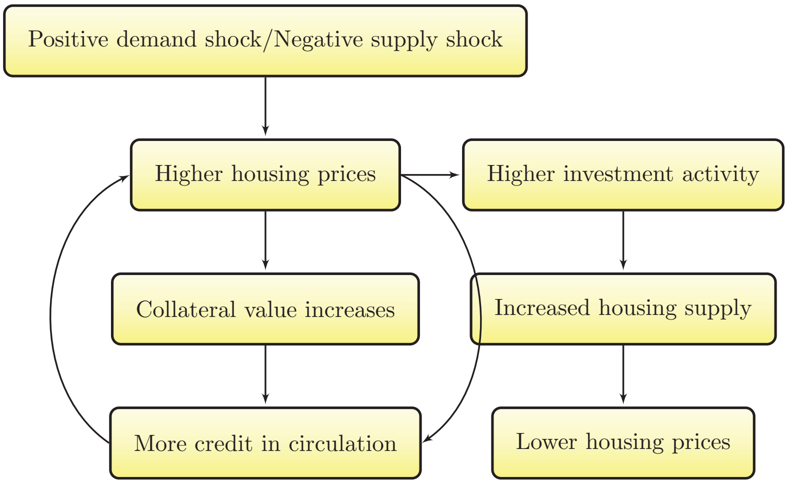


FIGURE 1.5: Two-way-interaction between housing prices and credit with housing supply side

It is the possible existence of a credit-housing price spiral that motivates the analysis in Chapter 2 of this thesis. In that chapter, the interaction between housing prices and household borrowing in Norway is investigated. The analysis shows that there exists strong evidence for the existence of a self-reinforcing feedback mechanism between housing prices and credit in Norway, both in the short-run and in the long-run. The analysis also reveals an important short-run effect from households' expectations about the development in their own and in the Norwegian economy. A similar feedback mechanism between the housing and the credit market has been documented for the case of Ireland, Greece and the US, see Fitzpatrick and McQuinn (2007); Brissimis and Vlassopoulos (2009); Berlinghieri (2010), respectively.

The econometric model for the joint determination of housing prices and credit that is presented in Chapter 2 is also extended to include a model for the supply side of the housing market, where housing investments (or – more precisely – housing starts) are modeled using an equation similar to (1.7). Theoretically, this is expected to dampen the housing-credit spiral, as illustrated in Figure 1.5. In that figure, we see how an increase in housing prices will result in an increased investment activity through a Tobin-Q effect, which in turn is manifested in an increased supply of dwellings. Naturally, the increase in supply will dampen the pressure on housing prices and therefore the entire financial accelerator effect.

Indeed, when incorporating the supply side into the simultaneous housing-credit system, the long-run effects on real housing prices following a shock to real disposable income or the real interest rate are substantially lowered, while the short-run effects are almost unaltered. The latter is due to lags in the construction process. The results presented in Chapter 2 suggests at least two ways in which a pressure on housing prices may be dampened. First, since housing prices are responsive to the supply of housing, the results suggest that measures limiting regulations on housing construction may be an effective tool to dampen the effect on housing prices following a demand shock. This is of course only true to the extent that relaxing such regulations contribute to increase the supply elasticity – an empirical question that is addressed using disaggregate US housing price data in Chapter 4 of this thesis. Second, the results indicate that constraints on banks' lending behavior may provide an effective tool to dampen excessive fluctuations in the housing market.

The relevance of developing econometric models of this kind has recently been exemplified in a report from Statistics Norway (see SSB (2013)).[8] In that report, the model for the joint determination of housing prices and credit that is presented in Chapter 2 of this thesis was successfully implemented in the operational macroeconometric model, KVARTS (Eika and Moum, 2005). KVARTS also includes a feedback from housing prices to consumption through wealth effects (see Jansen (2013)) and is therefore well suited for analyzing the real economic consequences of a housing-credit spiral.

While the credit growth in Norway has averaged 6–7 percent on annual basis in recent years, it was assumed in the simulations underlying the analysis in SSB (2013) that new capital requirements for the banks would decrease the growth in household debt by 0.5 percentage points in each of the quarters between 2013–2016. While this amounts to a reduction of about 2 percentage points at an annual basis in the absence of a housing-credit spiral, it was found that the annual growth in 2016 would be down by 2.9 percentage points due to the dampening effects this reduction in credit supply has on housing prices, and thereby on household debt. Relative to a reference path without any tightening of lending standards, the simulations suggest that these measures will lower housing prices by 7.2 percent by 2016, which again feeds into the real economy through both consumption wealth effects and through a lower activity in the construction sector. At the end of the simulation period, aggregate investments and private consumption are down by 1.6 percent and 1 percent relative to the reference paths, respectively.

1.2.1 Housing price expectations

It is evident from (1.2) that an important theoretical element of housing price determination is the expectation about future housing price gains, which affects housing prices by altering the user cost of housing. An assumption I make throughout this thesis is

[8] The report may be downloaded from http://www.ssb.no/en/nasjonalregnskap-og-konjunkturer/artikler-og-publikasjoner/_attachment/110907?_ts=13e5add4a20.

that price expectations may be modeled by lagged appreciation, i.e. that these expectations are formed adaptively. A similar approach has been followed by Muellbauer and co-authors who include a moving average of past housing price appreciation in the user cost term. Both approaches are consistent with the view in Abraham and Hendershott (1996), who interpret lagged housing price appreciation as capturing a bubble builder – or a momentum – effect, but the assumption that housing price expectations are formed adaptively rather than rationally calls for some justification given the strong position that rational expectations have in modern macroeconomics.

Perhaps surprisingly, there is strong evidence in the literature that housing price expectations are formed in an adaptive manner, see e.g. Jurgilas and Lansing (2013) and the references therein. In particular, survey evidence from the US for the years 2006 and 2007 (Shiller (2008)) suggests that individuals in areas with increasing housing prices expected further increases, while the opposite was the case in areas with recent declines in home values. Strikingly, conducting a similar survey in the midst of the national housing bust (in the year 2008), Case and Shiller (2012) find that individuals living in previously booming areas now expected a decline in housing prices.

To shed some more light on this assumption, I have collected quarterly survey data for Norway on the number of households expecting an increase and a decrease in housing prices over the next twelve months. The sample is relatively short and covers the period 2007q2–2011q4.[9] To investigate the role of past housing price appreciation on the net number of survey respondents believing in an increase in housing prices over the next year, I estimated a simple model of the following form by OLS:

$$E_t\left(\Delta ph_{t+1}\right) = \beta_0 + \sum_{i=0}^{4} \beta_{1+i} \Delta ph_{t-i} + u_t \qquad (1.10)$$

with $E_t\left(\Delta ph_{t+1}\right)$ denoting the net number of respondents expecting an increase in housing prices over the next 12 months and Δph_{t-i} measuring the quarterly price increase from period $t-i-1$ to period $t-i$. Results are summarized in Table 1.1. It is evident that lagged housing price appreciation does a fairly good job in explaining the net number of respondents expecting an increase in housing prices over the coming year with an adjusted R^2 of around 0.80. Acknowledging that the sample size is extremely short, it is still noticeable that the findings here are in line with the existing evidence in the literature, giving credence to the assumption that an adaptive expectation channel may be of relevance – at least in a housing market context.

[9] I am grateful to Kevin Lansing for sharing these data.

TABLE 1.1: The role of past housing price appreciation for future expectations

Variable	Coefficint	t-value
Constant	0.162	3.82
Δph_t	4.581	3.15
Δph_{t-1}	3.066	1.92
Δph_{t-2}	5.273	2.92
Δph_{t-3}	1.654	1.02
Δph_{t-4}	3.363	2.31
$Adj.R^2$	0.771	
Number of Observations	16	

Notes: This table reports the estimates obtained when I estimate equation (1.10) by OLS. The reported t-values are in absolute terms.

1.3 The role of fundamentals in housing price determination

While a credit-housing spiral is consistent with the existence of a bubble in the housing market, it need not imply so in any way. To discuss whether the development in any given housing market is best characterized as exercising bubble behavior, at least two requirements must be satisfied: first, we must have a conceptual understanding of what we define as a housing bubble. Second, given our conceptual understanding of a housing bubble, we need to have a formal (statistical) framework in which the existence of a bubble may be detected.

In that respect, there are two interesting observations that can be made from the alternative operationalizations of the theory model outlined in Section 1.1. A conventional metrics used by many institutions is to regard the *price-to-rent ratio* (the return to housing investments), or the *price-to-income ratio* (the affordability of housing), as measures of the temperature in the housing market. Yet another approach is followed in Cardarelli and Rebucci (2008), who uses the residuals from a model for housing price growth with price divided by per capita disposable income, interest rates, income growth and credit growth among the explanatory variables to define a *housing price gap* for a set of OECD countries over the 1997–2007 period. Common to these three approaches – and at the odds with the discussion in Section 1.1 – is that they miss important theoretical aspects like the user cost of housing and the supply of dwellings.[10] As pointed out by Muellbauer (2012), there is also a poor relationship between the housing price gaps estimated by Cardarelli and Rebucci (2008) and the subsequent busts in housing prices – with e.g. the

[10] The approach followed by Cardarelli and Rebucci (2008) is not subject to the critique of omitting the user cost to the same extent as the other two approaches, since they include measures of the interest rate, which clearly is an important component of the user cost. They do, however, remain subject to the critique of omitting information about the supply side.

estimated gap for the US being among the smallest. Muellbauer attribute this partly to the lack of a clear theoretical foundation – consistent with the above discussion. An evaluation of the temperature in the housing market may be best founded in a model that incorporates important theoretical drivers, such as the user cost of housing – and in case an inverted demand approach is pursued – the supply of dwellings.

Muellbauer and Murphy (2008) highlights that one way of detecting overheating in the housing market is by building an econometric model that links housing prices to the development in underlying economic fundamentals by use of historical data, and then investigate whether there are evidence of large deviations between actual housing prices and the value implied by these economic fundamentals. Large and persistent deviations of actual housing prices from the value implied by the economic fundamentals would then indicate an unsustainable development in housing prices. To illustrate how this may be implemented in practice, I have estimated a simple model of the form (1.5) by use of OLS over the period 1980q1–1995q4 on aggregate US data. The results are reported in Table 1.2.

Table 1.2: An estimated price-to-rent model for the US, 1980q1–1995q4

Variable	Coefficient	t-value
Constant	-4.22	24.9
UC	-0.93	14.7
r	0.96	26.3
σ	0.012	
Number of observations	64	

Notes: This table reports the estimates obtained when I estimate equation (1.5) by OLS using aggregate US data for the period 1980q1–1995q4. The reported t-values are measured in absolute value.

The first thing to notice is that the results are theoretically consistent, and that the (log of the) price-to-rent ratio is inversely proportional to the user cost. It is also re-assuring that the estimated coefficients are similar to those reported in Chapter 3 of this thesis, where I consider more sophisticated econometric models to look at the relationship between housing prices, rents and the user cost.

To investigate what this exercise may tell us about whether or not US housing prices were systematically overvalued at any point during the period 1980q1–1995q4, Figure 1.6 plots the actual housing price development over that period together with the fundamental value implied by the estimated model. The figure also displays the equilibrium deviations, i.e. the difference between the actual housing prices and the value implied by the simple econometric model. It is clearly seen that while there are periods of disequilibrium (actual prices not equal to model implied fundamental prices), there is also a tendency that housing prices return to the value implied by the fundamentals. Hence, judged by this measure, we would say that there are no clear signs of systematic overvaluation of US housing prices over this period, i.e. there is no evidence of bubble behavior.

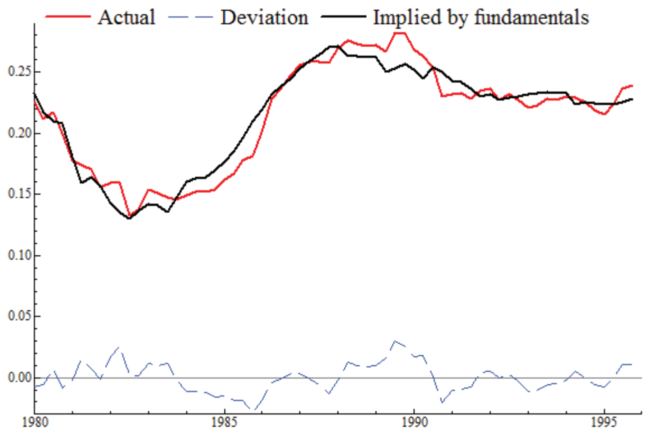


Figure 1.6: US housing prices and fundamentals, 1980q1–1995q4

Having estimated a model linking US housing prices to economic fundamentals, we can use the same model to ask whether there are evidence of systematic disequilibrium behavior during the period 1995q4–2006q4 – a period with a far more rapid price increase in aggregate US housing prices. For this purpose, I have used the estimates reported in Table 1.2 to construct a time series for the "model implied fundamental housing prices" over this period. This series is plotted together with the actual price development in Figure 1.7. Again, the figure also graphs the difference between actual housing prices and the value implied by the fundamentals, i.e. the equilibrium deviations.

Comparing Figure 1.7 to Figure 1.6, we see that the model tells a completely different story for this period! It is evident that, starting in the late 1990s/early 2000s, there was a growing gap between the actual housing price development and what was implied by the development of the underlying fundamentals. By the early the 2000s, this gap grows bigger and by 2006, only 50 percent of the housing price level may be attributed to the underlying fundamentals.

This simple – and easily implementable – analysis suggest that there was a growing disconnect between housing prices and fundamentals in the US housing market in the 2000s. The topic of Chapter 3 of this thesis is to investigate this disequilibrium behavior in more detail. In that chapter, I ask two key questions: could we by the aid of real time econometric modeling have detected these imbalances in real time, and what

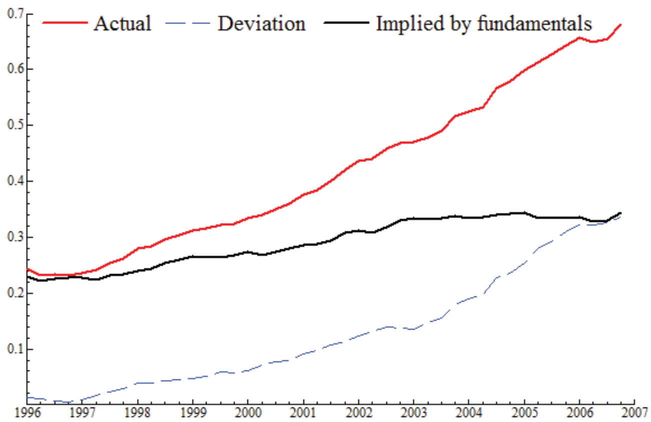


FIGURE 1.7: US housing prices and fundamentals, 1995q4–2006q4

factors may explain the increasing disconnect between US housing prices and fundamentals in the early 2000s? To answer these questions, I take as a starting point the two operationalizations of the life-cycle model, as represented by (1.5) and (1.6).

Since all variables in these equations are found to exhibit stochastic non-stationarities (of first order), the question of whether housing prices are determined by fundamentals boils down to a question of whether there exists evidence of cointegration between housing prices and these non-stationary economic variables, i.e. whether it can be established that $ph - \beta_r r - \beta_{UC}UC \sim I(0)$ and $ph - \beta_y y + \beta_h h - \beta_{UC}UC \sim I(0)$. An additional requirement I impose for the detection of bubble behavior is that cointegration can be established on a given sample ($t = 1, \ldots, T_1$), with stable coefficients, while disappearing when the bubble period is included in the sample ($t = t = 1, \ldots, T_1, T_1 + 1, \ldots T$, with the bubble period running from $T_1 + 1$ to T).

My results do indeed indicate that the imbalances in the US housing market could have been detected with the aid of real time econometric modeling. I take the analysis a step further and develop two "bubble indicators" that both clearly suggest a bubble in US housing prices at a quite early stage. Taking the analysis yet another step further, I find that the US housing bubble – as detected by these indicators – may be attributed to increased borrowing to the subprime segment. This is a different explanation than what has been put forth by Caballero and Krishnamurthy (2009) and Taylor (2008, 2009), who ascribe the housing market imbalances to the large capital inflows and loose monetary policy. It is, however, in accordance with the conclusions of Chapter 4 of this thesis, where it is documented that differences in the exposure to aggressive lending products is important in explaining regional differences in the cumulative housing price growth over the period 2000–2006 for US metro areas. The result is also in line with Duca et al. (2011a,b), who find that accounting for exogenous shifts in credit standards – as measured by the loan-to-value ratio for first time buyers – is important to build a reliable econometric model for aggregate US housing determination.

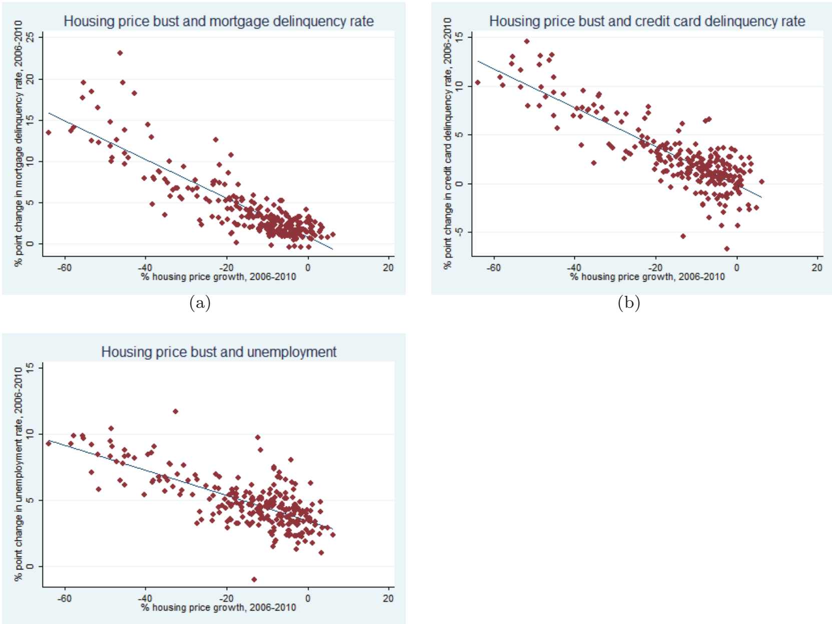
1.4 Spatial differences

The international housing price boom that started in the mid 1990s (confer Figure 1.1) was recognized by an increased synchronization of housing price movements across countries (Kim and Renaud, 2009 and Girouard et al. (2006)). That said, Figure 1.1 demonstrates that there were substantial cross country variations as well. Also at a subnational level there exists enormous differences across geographical areas. For the case of the US, possible explanations of these differences are addressed in e.g. Glaeser et al. (2008), Huang and Tang (2012) and in Chapter 4 and Chapter 5 of this thesis.

As Figure 1.8 demonstrates, the Metropolitan Statistical Areas that experienced the greatest drop in real house prices in the 2006-2010 period are the same areas that had the greatest increase in unemployment and delinquency rates on mortgages and credit card loans over that period.

Mian and Sufi (2010) show that the areas which experienced the greatest run-ups in household leverage are the same areas that saw the greatest fall in consumption and the greatest hike in unemployment rates. At the same time, Mian and Sufi (2009) and Pavlov and Wachter (2011) have shown that areas with more subprime lending also witnessed a greater build-up of housing prices, while Goetzmann et al. (2012) have shown a positive impact of housing price appreciation on approval rates. Thus, given the close interconnection between lending standards, housing prices and the real economy, understanding what determines the cross sectional variation in housing price volatility seems to be a particularly relevant issue.

Regional differences in housing price developments are due to both demand and supply factors. A factor that may be especially important in this respect is differences in the supply elasticity. Areas with an inelastic housing supply will have a greater price increase following a demand shock. How responsive supply is to an increase in prices may depend on several factors, and in particular land availability constraints and regulatory constraints on housing construction. The importance of the supply elasticity following a negative demand shock is however less clear, due to the durability of housing and the fact that supply is usually rigid downwards.



(a) (b) (c)

FIGURE 1.8: Housing prices during the boom vs. delinquency rates and the unemployment rate during the bust for 247 US MSAs

As a first look at the relationship between the volatility in housing prices over the recent US boom-bust cycle and the degree of supply restrictions, as well as subprime exposure, I ask the question of why a given area experienced an above average housing price increase (decrease) – a super boom (super bust). The dependent variable in both specifications is a dummy variable taking the value one if an area experienced a housing price growth that (in absolute value) exceeded the average price growth of the 247 US Metropolitan Statistical Areas included in the sample. For the boom period, which I take as the period 2000–2006, I estimate an equation of the following form:

$$\begin{aligned} Superboom_i = \mu^{Boom} &+ \beta^{Boom}_{wrluri} wrluri_i + \beta^{Boom}_{unaval} unaval_i \\ &+ \beta^{Boom}_{\Delta y} \Delta y^{Boom}_i + \beta^{Boom}_{credit} credit^{Boom}_i + \varepsilon^{Boom}_i \end{aligned} \tag{1.11}$$

where $wrluri$ is the regulatory supply restriction index developed by Gyourko et al. (2008), $unaval$ is the geographical supply restriction index of Saiz (2010)[11], Δy^{Boom} is the percentage change in personal income during the boom, and $credit^{Boom}$ is the cumulative increase in subprime lending per capita over the same period. The specification is estimating using a probit specification, and results are recorded in Table 1.3.

[11] Higher values of the indexes indicate a more restricted supply.

TABLE 1.3: What explains the super booms?

Variable	Coefficient	t-value
$Constant$	3.47	1.50
$wrluri$	4.29	4.25
$unaval$	4.22	6.42
Δy_t^{Boom}	4.72	3.87
$credit_t^{Boom}$	0.48	2.95
Number of observations	247	

Notes: This table reports the estimates obtained when I estimate equation (1.11) using a probit specification and data for 247 US Metropolitan Statistical areas. The boom is taken to run through the period 2000–2006, and the reported t-values are measured in absolute value.

While these results are only indicative due to the quasi-reduced form nature of the specification[12], they still tell an interesting story. As would be expected, areas that had a stronger income growth or a greater exposure to subprime lending during the boom were more likely to experience an above average increase in housing prices. Further, we see that both regulatory and geographical supply restrictions affect the probability of an area experiencing a super boom positively, i.e. this simple modeling exercise suggests that areas *with* many regulatory an physical supply restrictions were more likely to experience a greater housing price boom than areas *without* such restrictions.

To explore the relevance of such restrictions in explaining the price drop during the 2006–2010 bust period, I estimate a model of the following form:

$$\begin{aligned} Superbust_i = \mu^{Bust} &+ \beta_{wrluri}^{Bust} wrluri_i + \beta_{unaval}^{Bust} unaval_i \\ &+ \beta_{\Delta y}^{Bust} \Delta y_i^{Bust} + \beta_{credit}^{Bust} credit_i^{Boom} + \varepsilon_i^{Bust} \end{aligned} \tag{1.12}$$

with Δy^{Bust} measuring the income growth over the 2006–2010 period, while all other variable definitions are as defined previously. Results are displayed in Table 1.4.

It is clearly seen that areas with a higher income growth during the 2006–2010 period had a lower probability of experiencing a super bust, which is in accordance with what we would expect from a theoretical point of view. In addition, the results suggest that areas that were more exposed to subprime lending during the boom period had a greater probability of experiencing a super bust, i.e. the more aggressive the lending during the boom, the greater the price drop during the bust. Furthermore, we see that both measures of supply restrictions affect the probability of an area experiencing a

[12]By quasi-reduced form specification, I mean a specification that is neither reduced form nor part of a structural model. With reference to the discussion on Section 1.1, it is clear that the specification I consider here does not include any information about the supply side (except the regulatory measures of course).

TABLE 1.4: What explains the super busts?

Variable	Coefficient	t-value
Constant	10.30	4.16
wrluri	2.51	2.91
unaval	2.19	3.88
Δy_t^{Bust}	-7.76	3.80
$credit_t^{Boom}$	0.79	4.88
Number of observations	247	

Notes: This table reports the estimates obtained when I estimate equation (1.12) using a probit specification and data for 247 US Metropolitan Statistical areas. The bust is taken to run through the period 2006–2010, and the reported t-values are measured in absolute value.

super bust positively. This gives a first indication that regulatory and physical supply restrictions not only leads to a greater boom, but that they also magnify the size of the housing bust.

While the results from these simple models are interesting in their own right, they are silent about what mechanisms causes an area to experience a super bust. In addition, they do not account for the potential simultaneity between subprime lending and housing prices, and we are not able to distinguish between the effects such supply restrictions have on housing prices and housing supply, respectively. That said, we get a first indication that two important factors in explaining the regional differences in housing price volatility during the recent US boom-bust cycle is the exposure to aggressive lending products and differences in restrictions on housing supply.

Chapter 4 of this thesis investigates these issues in more detail by estimating a fully simultaneous equation system consisting of equations for housing prices (an inverted demand equation), quantity (a housing supply equation) and subprime lending. There are two main innovations in the analysis of that chapter: first, we allow for a financial accelerator effect by letting prices depend on subprime lending, and *vice versa*. Second, we allow supply side restrictions to affect the dynamics of the housing market by altering the supply elasticity. A clear advantage of this modeling approach is that we can distinguish between the price and quantity response following a positive demand shock for areas with different degrees of supply restrictions. Further, we can investigate how the importance of a financial accelerator effect differs along the same dimension.

The analysis leads to several interesting conclusions. First, in line with the results reported in Table 1.3, we find that areas with many restrictions on the supply side have a greater price reaction following a positive demand shock. Areas with fewer restrictions on the supply side absorb most of the shock in terms of quantity adjustments. That said, the price increase sets in motion a financial accelerator mechanism, where higher housing prices leads to more subprime lending, which again puts an upward pressure on prices. Consequentially, supply will increase as well. This self-reinforcing feedback mechanism is stronger in areas with a lower supply elasticity, since the initial price reaction is greater in these areas. The end result is that even though some areas have many supply restrictions – affecting the supply elasticity negatively – the total increase in quantity following a demand shock is almost the same independent of the degree of supply side restrictions. Thus, while we find an unambigiously greater price response in more regulated areas, it is not clear that the quantity increase will be greater in less restricted areas once the financial accelerator is taken into account.

These findings have interesting implications for the price drop during the bust, since both the price and the quantity overhang will tend to have a negative impact on housing prices when the demand shock is reversed. Since the price overhang is markedly greater in the restricted areas, while there is no big difference in the quantity overhang, we find that restricted areas are hit worse during the bust. Thus, the combination of a financial accelerator effect and supply restrictions documented in Chapter 4 provides one explanation to the results reported in Table 1.4.

The final chapter of this thesis – Chapter 5 – also pays heed to what factors contribute to explain the huge cross sectional variations in local US housing prices, and in particular why some areas experienced a greater housing boom than others. Supporting the results in Chapter 4, we find that areas with a low supply elasticity were more affected by subprime lending. Furthermore, we find that an adaptive expectations channel is more important in areas with many such restrictions.

Though the price development across regional markets may well differ – especially in the short and medium term – Meen (1999, 2001) and Holmes et al. (2011) points to several channels which may cause prices to converge across different areas, i.e. a ripple effect. Four channels that are mentioned as a possible explanation of why a ripple effect could occur are migration, equity transfer, spatial arbitrage and spatial patterns in the determinants of house prices (see Meen (1999) for more discussion). That said, there may exist frictions that pull in the other direction, i.e. a sustained divergence of prices across areas may occur (see e.g. the discussion in Meen (2001)). If a given market is very distant from other markets, the search related costs will be very high as well, which suggests that a price differential may be sustained.

Following Meen (1999, 2001), there has been several discussions in the literature on the importance of ripple effects between housing markets. While many papers have investigated this empirically on UK data (see e.g. Holmes and Grimes (2008) and Cameron et al. (2005) and the references therein), the literature on US data is relatively scarce.

Gupta and Miller (2012) consider the ripple effect between Los Angeles, Las Vegas and Phoenix employing quarterly data for the period 1978-2008. Using the Johansen (1988) procedure, they find that there exists one cointegrating vector between the three price indexes. It is found that prices in Los Angeles Granger causes prices in Las Vegas, while prices in Las Vegas Granger causes prices in Phoenix. Other than that, there is no evidence of Granger causality. The authors interpret this as evidence of a ripple going from Los Angeles to Las Vegas, and then to Phoenix.

In line with this, Holmes et al. (2011) use the housing price differential across areas as an indicator for regional housing price convergence. In particular, using the pair-wise procedure suggested by Pesaran (2007) and Pesaran et al. (2009) for 48 US states over the period 1975q1–2008q4, they find evidence of a regional housing price convergence. Furthermore, they find that the distance between regions is important in explaining this convergence.

Another way of analyzing ripple effects is by use of a spatial VAR (SpVAR) model. This approach has been adopted by Kuethe and Pede (2010), who looks at the spill-over effects between the eleven US states belonging to the Western region by considering a sample spanning the period 1988q1–2007q4. To connect the different areas in the spatial domain, the authors assign a value one to areas that are bordering the area under consideration, and zero otherwise. Based on this, they construct a weighting matrix that links the different areas together. Tests for Granger non-causality show that there is evidence of a spatial spill-over across states within the Western region. Interestingly, and contradicting Vansteenkiste (2007), Kuethe and Pede (2010) find that California is particularly affected by its neighbors.

A third empirical methodology that has recently been applied to analyze the ripple effect is to consider a global vector autoregressive model (GVAR), see Pesaran et al. (2004), Dees, di Mauro, Pesaran, and Smith (2007) and Dees, Holly, Pesaran, and Smith (2007) for details on the GVAR model. Vansteenkiste (2007) and Vansteenkiste and Hiebert (2011) aim to explore the linkages between regional housing markets using this approach. While the first paper focus on state level spill-overs in the US housing market, the second paper considers similar spill-overs for 7 Euro-area countries. Both papers use the same information set, i.e. housing prices, income and an interest rate variable. While Vansteenkiste (2007) finds evidence of spill-overs in the US, there is less evidence of such an interconnection for the Euro-area countries considered by Vansteenkiste and Hiebert (2011). Vansteenkiste (2007) further finds that California exercises the greatest influence on other regional housing markets. Contrary to what would be expected based on the results documented in Chapter 4, the author finds that an interest rate shock has stronger effects in a relatively supply elastic state such as Texas than in more restricted areas such as California and Florida. A weakness of these papers is that no measure of the supply side is included in the econometric models of housing price formation, i.e. the specifications are quasi-reduced form by nature.

While none of the chapters in this thesis are explicitly concerned with investigating the relevance of the ripple effect, it is clearly important to understand the possible existence of contagion effects across regional housing markets. The results developed in Chapter 5 of this thesis may however be an important first step in that respect, since separate econometric models are developed for the 100 largest Metropolitan Statistical Areas in the US. These models may form the basis for a GVAR model that can be used to analyze the relevance of the ripple across local US housing markets. That is the topic of an ongoing research project with Christian H. Christensen, where we aim at consolidating the results reported in Chapter 5 with a spatial spill-over using the GVAR methodology. If successful, this project will give the basis for analyzing how a shock in e.g. San-Francisco may affect the housing price development in a nearby area such as San-Diego, or even a distant area such as Boston, through a ripple effect.

A summary of Chapter 2–5 of the thesis follows below.

Chapter 2: Self-reinforcing effects between housing prices and credit

Chapter 2 of this thesis is a result of joint work with Eilev S. Jansen, and a shorter version of the chapter has been published in *Journal of Housing Economics*, see Anundsen and Jansen (2013a). However, in revising the paper for journal publication, we were asked to move some technical details and results to an online appendix. While these details are available on my personal webpage, http://www.andre-anundsen.com/, we have also published an extended version of the paper that includes the parts we omitted in the journal article as a Discussion Paper (see Anundsen and Jansen (2013b)). In addition, in the extended version we considered a few extensions and robustness checks of the model. For completeness, it is the extended version of the paper (i.e. Anundsen and Jansen (2013b)) that is included as Chapter 2 of this thesis.

The aim of Chapter 2 is to investigate whether there exists a self-reinforcing feedback mechanism between housing prices and household borrowing in Norway. The system based cointegration approach of Johansen (1988) is used to explore whether there is evidence of a housing-credit spiral in the long-run, while the short-run dependence is investigated by estimating a fully structural vector equilibrium correction model (SVECM) by full information maximum likelihood techniques. The sample runs through the period 1986q2–2008q4, i.e. the period after which the Norwegian housing and credit markets can be considered fully deregulated.

The results from the cointegration analysis suggest that there exists a long-run two-way interaction between housing prices and household borrowing in Norway, giving rise to a financial accelerator effect. The long-run results also suggest an important role of monetary policy in containing excessive housing price increases by dampening the the amount of credit in circulation through interest rate adjustments. The short-run housing-credit system is built around the two cointegrating vectors we find in the cointegration analysis. The dynamic models suggest an important role of expectations about the future economic development in driving housing prices in the short-run. Using the model to conduct various simulations, we show that there is indeed a financial accelerator at work in the Norwegian housing and credit markets.

We consider an extended version of the housing-credit system by incorporating a small model for the supply side of the Norwegian housing market that takes into account that higher housing prices stimulates increased construction activity. Simulations from the extended model demonstrates that the increased construction activity mainly contributes to dampen the long-run effects on housing prices and credit following a given demand shock, since it takes time before new housing starts are turned into actual dwellings.

Part of the evaluation of the model also concerns its out-of-sample forecasting performance. It is shown that the model does a fairly good job in forecasting both housing prices and credit over the 2009–2011 period. However, there are some significant forecast errors for credit in 2010q1 and 2011q1. This may be attributed to the very cold winters in Norway during those years, which made the consumption deflator used for the nominal-to-real transformations jump upwards. Indeed, when we allow for short-run inflation effects, the forecasting properties of the model is improved significantly. The evaluation of the model on an extended sample demonstrates the recursive stability of the coefficients in the model and shows that it is producing satisfying forecasts also for all quarters in 2012.

In an ongoing research project, we evaluate the *ex ante* forecasting properties of the model against a range of alternative forecasting models, and preliminary results are promising.

Chapter 3: Econometric regime shifts and the US subprime bubble

Chapter 3 of this thesis has been accepted for publication in *Journal of Applied Econometrics*, see Anundsen (2013). The chapter does, however, include some minor reporting additions that I think are useful for documentation purposes.

The contributions of this chapter are threefold. First, there has been a long standing debate in the academic literature on the role of so called economic fundamentals in US housing price determination, see e.g. Abraham and Hendershott (1996); Malpezzi (1999); Meen (2002); McCarthy and Peach (2004); Gallin (2006, 2008); Mikhed and Zemcik (2009a,b); Zhou (2010); Clark and Coggin (2011); Duca et al. (2011a,b). While half of these studies have found that US housing prices may be explained by underlying economic fundamentals, the other half has concluded opposite. It is shown in Chapter 3 – using data for the period 1975q1–2010q4 – that the conflicting results may be attributed to an econometric regime shift in two alternative models of US housing price formation in the early 2000s, i.e. considering different sample end points, I am able to encompass the previous findings. To arrive at these results, I make use of both the system based cointegration approach due to Johansen (1988) and a conditional equilibrium correction model approach.

Second, the role of fundamentals is also important in determining whether or not imbalances are building up. I therefore develop two econometrically based "bubble indicators" that are based on the premise that there is a bubble whenever an econometric model that links housing prices to a set of economic fundamentals – that for previous periods have been shown to yield meaningful results and stable coefficients – breaks down. The indicators clearly demonstrate that the US housing market experienced bubble behavior already in the early 2000s. Furthermore, these indicators are shown to Granger cause a set of financial crisis related measures.

The final contribution of this chapter is concerned with possible explanations of the bubble behavior – as detected by the bubble indicators – of US housing prices in the 2000s. Caballero and Krishnamurthy (2009) and Taylor (2008, 2009) have ascribed the housing market imbalances to the large capital inflows and loose monetary policy, respectively. My results suggest a different explanation. In particular, I find that the econometric models that break down in the early 2000s – interpreted as a bubble – do not break down once I control for the relaxation of credit market constraints, as measured by the exposure to subprime lending. This suggests that the bubble was caused by the increased borrowing to the subprime segment.

Chapter 4: Supply restrictions, subprime lending and regional US housing prices

Chapter 4 of this thesis is the product of joint work with Christian H. Christensen. The chapter is concerned with how the interaction between housing supply restrictions and mortgage credit constraints affects housing price volatility. The paper is currently under review in *Journal of Money Credit and Banking* that will publish a special issue in relation to the conference “Housing, Stability and the Macro Economy: International Perspectives” – at which the paper was included in the conference program.

Although the boom and the bust in US housing prices was substantial at a national scale, an aggregate analysis hides the enormous differences at the subnational level. There are many reasons why the price development may differ along the geographical dimension. First, the development on the demand side will differ, with some cities experiencing a more fortunate development in e.g. income. In addition, differences on the supply side may be substantial. Gyourko et al. (2008) and Saiz (2010) have developed subnational measures for the degree of politically enforced and geographically determined restrictions on land supply, which are sought to affect the supply elasticity and are shown to vary greatly across US cities.

Chapter 4 analyzes the role of supply side restrictions – as measured by the indexes alluded to above – in explaining the enormous regional differences in housing price volatility over the recent boom and bust cycle in the US. In addition to focusing on the importance of supply restrictions in local housing price determination, we also pay attention to regional differences in the relaxation of credit constraints – as measured by subprime exposure – during the recent housing boom as a possible source of explaining the huge heterogeneity that is apparent in the data.

For this purpose, we exploit data for 247 Metropolitan Statistical Areas to identify a supply-demand system using full information maximum likelihood methods. In addition to identifying an inverted demand equation (normalized with respect to prices), we also identify a supply equation (normalized with respect to quantity) where the supply elasticity is directly linked to the degree of geographical and political supply restrictions. Numerous papers (see e.g. Titman (1985); Mayer and Somerville (2000); Malpezzi and Maclennan (2001); Green et al. (2005); Saiz (2010)) agree that these restrictions should be important for the supply elasticity, but our paper is – to the best of our knowledge – the first to identify a structural supply equation where the supply elasticity is directly linked to these measures. This has several interesting implications. First, it allows us to distinguish between the quantity and price reactions following a demand shock in areas with different degrees of supply restrictions. In accordance with the theory, we find that supply restricted areas will witness a larger price increase following a positive demand shock, while more of the shock is absorbed in terms of an increase in quantity in less restricted areas. The first finding corroborates the reduced form results of Glaeser et al. (2008) and Huang and Tang (2012), but an advantage of our analysis is that we can attribute this to a lower supply elasticity in these areas, since we find that the supply restrictions affect the elasticity of supply negatively.

While this is an interesting finding, we go further and augment our supply-demand system with an equation for subprime lending, where we find that subprime lending depends positively on housing prices, and *vice versa*. This gives rise to an endogenous feedback mechanism between housing prices and subprime lending, with interesting implications for both the boom and bust period price determination. First, since supply restricted areas experience a greater price increase following an increase in demand, these areas will also have a stronger price-subprime spiral, which pushes prices even further up – with the implication that a positive demand shock leads to a substantially larger price increase in supply restricted areas. Second, since prices increase much more in these areas, the total quantity response following a demand shock is almost independent of the supply elasticity. This is because the supply response depends both on how much prices increase and on the supply elasticity (confer (1.7) above). While the unrestricted areas see a larger increase in quantity for a given change in prices, the total change in prices is much higher in the restricted areas, which almost outweighs this effect. This finding is consistent with the patterns observed in the data, where there is no clear relationship between the two supply restriction indexes and the total quantity change over the 2000–2006 boom period.

Incorporating a financial accelerator into the model also have interesting implications for the bust period price dynamics. Since the supply of housing is downward rigid, it can be shown that the bust period price response – interpreted as the effect of reverting the positive demand shock – will depend negatively on both the price and the quantity response during the boom. Since we find that the quantity response is almost disconnected with the supply elasticity, while the price response is higher the lower is the elasticity of supply, it follows that restricted areas are predicted to experience a greater drop in prices during the bust. While this finding is in accordance with the reduced form results of Huang and Tang (2012), the modeling setup pursued in our paper allow us to interpret this as partly caused by the existence of an endogenous feedback mechanism between housing prices and subprime lending that is stronger the more restricted is the supply of dwellings.

Chapter 5: Regional US housing price formation: Does one size fit all?

The final chapter of this thesis, Chapter 5, is also a result of my collaboration with Christian H. Christensen, and – like Chapter 4 – this chapter is concerned with regional differences in US housing price determination. That said, some of the key questions we ask and – in particular – the methodological approach adopted differ substantially from Chapter 4.

We ask three key questions that are relevant for the modeling of regional US housing prices. First, we ask whether there are signs of coefficient heterogeneity in regional long-run housing price determination, i.e. whether a "one size fits all" approach to modeling regional US housing prices works well or not. This is important both for the choice of econometric model and to get a proper understanding of local US housing price determination. Second, we ask whether the role of subprime lending during the recent housing boom was different across regional housing markets. Finally, having established that there are major heterogeneities in regional housing price determination, we investigate whether time invariant and regional specific factors may explain the coefficient heterogeneity and the differences in the importance of subprime lending.

Exploiting a panel data set covering the 100 largest metropolitan statistical areas in the US over the period 1980q1–2010q2, we start by estimating an inverted demand equation by use of several different econometric techniques allowing for different degrees of coefficient heterogeneity. Tests for poolability (slope homogeneity) show that the hypothesis of equal long-run coefficients is firmly rejected. This clearly demonstrates that separate regional models are needed to understand local US housing price formation. For that reason, we develop separate cointegrated VAR models for all 100 areas. The results from the separate econometric models provide several interesting insights. In particular, we find that the role of subprime lending differed markedly across regional markets during the recent boom, and that there are important geographical differences in the importance of lagged housing price appreciation – which – as suggested by Muellbauer and Murphy (2008) – is a potential mechanism for overshooting. Abraham and Hendershott (1996) give a similar interpretation to the coefficients on lagged housing price appreciation by referring to it as a bubble builder – or a momentum – effect.

We take the analysis one step further, and ask what factors may explain the observed heterogeneity. For that purpose we exploit a set of cross sectional models and a logit specification. Our findings suggest that subprime lending was more important for housing price determination in areas with many restrictions on housing supply, as measured by the geographical restrictions index of Saiz (2010) and the regulatory supply restriction index of Gyourko et al. (2008). This result corroborates the findings presented in Chapter 4 of this thesis. In addition, we find evidence suggesting that an adaptive expectations channel – the "bubble builder" – is more important in areas with many restriction on land supply. The bubble builder effect is also found to be of a greater magnitude in more populous areas and in areas belonging to a state where lending is non-recourse. This may possibly be explained by herd behavior being more prevalent in bigger cities and that the perceived risk of a housing purchase is lower if lending is non-recourse. While our results indicate that the disequilibrium adjustments are restored more slowly in areas with non-recourse lending, we do not find a relationship between most of the long-run elasticities and the time invariant explanatory variables that we consider.

We hope that the models developed in Chapter 5 can be successfully implemented into a larger model allowing for spatial spill-overs across US metro areas – a project that is at our current research agenda.

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Appendix 1.A The life-cycle model of housing

To solve the life-cycle model, simply formulate the Hamiltonian in the following way:

$$\begin{aligned}\Lambda = e^{-\beta t}U(C_t, H_t) + \lambda_{1,t}e^{-\beta t}\left[(1-\theta_t)Y_t + (1-\theta_t)i_tW_t - PH_tI_t - S_t - C_t\right] \\ + \lambda_{2,t}e^{-\beta t}\left[I_t - \delta H_t - \dot{H}_t\right] + \lambda_{3,t}e^{-\beta t}\left[S_t - \pi_t W_t - \dot{W}_t\right]\end{aligned} \tag{1.A.1}$$

which yields the following first order conditions:

$$\frac{\partial \Lambda}{\partial C_t} : e^{-\beta t}U_C - \lambda_{1,t}e^{-\beta t} = 0 \Rightarrow U_C = \lambda_{1,t} \tag{1.A.2}$$

$$\frac{\partial \Lambda}{\partial I_t} : -\lambda_{1,t}e^{-\beta t}PH_t + \lambda_{2,t}e^{-\beta t} = 0 \Rightarrow PH_t = \frac{\lambda_{2,t}}{\lambda_{1,t}} \tag{1.A.3}$$

$$\frac{\partial \Lambda}{\partial S_t} : -\lambda_{1,t}e^{-\beta t} + \lambda_{3,t}e^{-\beta t} = 0 \Rightarrow \lambda_{1,t} = \lambda_{3,t} \tag{1.A.4}$$

$$\begin{aligned}\frac{\partial \Lambda}{\partial H_t} = \frac{d}{dt}\frac{\partial \Lambda}{\partial \dot{H}_t} &: e^{-\beta t}U_H - \lambda_{2,t}e^{-\beta t}\delta = \frac{d}{dt}\left(-\lambda_{2,t}e^{-\beta t}\right) \\ &\Rightarrow e^{-\beta t}U_H - \lambda_{2,t}e^{-\beta t}\delta = \lambda_{2,t}\beta e^{-\beta t} - \dot{\lambda}_{2,t}e^{-\beta t} \\ &\Rightarrow U_H - \lambda_{2,t}\delta = \lambda_{2,t}\beta - \dot{\lambda}_{2,t}\end{aligned} \tag{1.A.5}$$

$$\begin{aligned}\frac{\partial \Lambda}{\partial W_t} = \frac{d}{dt}\frac{\partial \lambda}{\partial \dot{W}_t} &: \lambda_{1,t}e^{-\beta t}(1-\theta_t)i_t - \lambda_{3,t}\pi_t e^{-\beta t} = \frac{d}{dt}\left(-\lambda_{3,t}e^{-\beta t}\right) \\ &\Rightarrow \lambda_{1,t}e^{-\beta t}(1-\theta_t)i_t - \lambda_{3,t}\pi_t e^{-\beta t} = \lambda_{3,t}\beta e^{-\beta t} - \dot{\lambda}_{3,t}e^{-\beta t} \\ &\Rightarrow \lambda_{1,t}(1-\theta_t)i_t - \lambda_{3,t}\pi_t = \lambda_{3,t}\beta - \dot{\lambda}_{3,t}\end{aligned} \tag{1.A.6}$$

Combining (1.A.4) and (1.A.6), we get:

$$\begin{aligned}\lambda_{1,t}\left[(1-\theta_t)i_t - \pi_t\right] = \beta\lambda_{1,t} - \dot{\lambda}_{1,t} \\ (1-\theta)i_t - \pi = \beta - \frac{\dot{\lambda}_{1,t}}{\lambda_{1,t}}\end{aligned} \tag{1.A.7}$$

Note that (1.A.5) may be rewritten in the following way

$$\frac{U_H}{\lambda_{2,t}} - \delta = \beta - \frac{\dot{\lambda}_{2,t}}{\lambda_{2,t}} \tag{1.A.8}$$

Combining (1.A.7) and (1.A.8), we find

$$(1-\theta_t)i_t - \pi_t + \frac{\dot{\lambda}_{1,t}}{\lambda_{1,t}} = \frac{U_H}{\lambda_{2,t}} - \delta + \frac{\dot{\lambda}_{2,t}}{\lambda_{2,t}}(1-\theta)i_t - \pi_t + \delta - \frac{U_H}{\lambda_{2,t}} = \frac{\dot{\lambda}_{2,t}}{\lambda_{2,t}} - \frac{\dot{\lambda}_{1,t}}{\lambda_{1,t}} \tag{1.A.9}$$

Considering (1.A.3), we see that:

$$\dot{PH}_t = \frac{\dot{\lambda}_{2,t}\lambda_{1,t} - \dot{\lambda}_{1,t}\lambda_{2,t}}{\lambda_{1,t}^2} \tag{1.A.10}$$

Now, divide (1.A.10) by (1.A.3) to get:

$$\begin{aligned}
\frac{\dot{PH}_t}{PH_t} &= \frac{\frac{\dot{\lambda}_{2,t}\lambda_{1,t} - \dot{\lambda}_{1,t}\lambda_{2,t}}{\lambda_{1,t}^2}}{\frac{\lambda_{2,t}}{\lambda_{1,t}}} = \frac{\lambda_{1,t}\left(\dot{\lambda}_{2,t}\lambda_{1,t} - \dot{\lambda}_{1,t}\lambda_{2,t}\right)}{\lambda_{2,t}\lambda_{1,t}^2} \\
&= \frac{\dot{\lambda}_{2,t}\lambda_{1,t} - \dot{\lambda}_{1,t}\lambda_{2,t}}{\lambda_{2,t}\lambda_{1,t}} \\
&= \frac{\dot{\lambda}_{2,t}}{\lambda_{2,t}} - \frac{\dot{\lambda}_{1,t}}{\lambda_{1,t}}
\end{aligned} \tag{1.A.11}$$

Inserting for (1.A.11) in (1.A.9), we get:

$$(1-\theta_t)i_t - \pi_t + \delta - \frac{U_H}{\lambda_{2,t}} = \frac{\dot{PH}_t}{PH_t} \tag{1.A.12}$$

Finally, inserting for $\lambda_{2,t} = \lambda_{1,t}PH_t$ from (1.A.3) in (1.A.12), while also remembering that $\lambda_{1,t} = U_C$ from (1.A.2), we obtain:

$$(1-\theta_t)i_t - \pi_t + \delta - \frac{U_H}{U_C PH_t} = \frac{\dot{PH}_t}{PH_t}$$

Rearranging slightly, we have:

$$\frac{U_H}{U_C} = PH_t\left[(1-\theta_t)i_t - \pi_t + \delta - \frac{\dot{PH}_t}{PH_t}\right] \tag{1.A.13}$$

which is the expression that forms the basis for many papers on housing price determination, i.e. the expression reported in (1.1).

Chapter 2

Self-reinforcing effects between housing prices and credit

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Abstract: The financial crisis has brought the interaction between housing prices and household borrowing into the limelight of the economic policy debate. This paper examines the nexus of housing prices and credit in Norway within a structural vector equilibrium correction model (SVECM) over the period 1986q2-2008q4. The results establish a two way interaction in the long-run, so that higher housing prices lead to a credit expansion, which in turn puts an upward pressure on prices. Interest rates influence housing prices indirectly through the credit channel. Furthermore, households' expectations about the future development of their own income as well as in the Norwegian economy have a significant impact on housing price growth. Dynamic simulations show how shocks are propagated and amplified. When we augment the model to include the supply side of the housing market, these effects are dampened.

Keywords: *Housing Prices; Household Borrowing; Financial Accelerator; SVECM; Dynamic Simulations*

JEL classification: *C32, C52, E27, E44, G21, G28, R21, R31*

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2.1 Introduction

The world wide financial crisis that originated with the US sub-prime crisis of 2007 has highlighted the importance of the interplay between financial markets and the real economy. A great number of factors contributed to the current crisis, see IMF (2009), Hubbard and Mayer (2009) and Acharia and Schnabl (2009). However, it seems to be widely agreed that it was primarily an unsustainable weakening of credit standards that induced the US mortgage lending and housing bubble. Countries with more stable credit conditions were mainly affected through the international financial linkages, e.g. European banks incurring heavy losses on securities tightly connected to the US mortgage market in the wake of the meltdown. In those countries, as Duca et al. (2010) emphasize, any overshooting of construction and housing prices owed more to traditional housing supply and demand factors.

However, there is a two-way direction of causation since imbalances in the housing market oftentimes have threatened the stability of the financial sector. In the past, there have been numerous episodes where falling housing prices have preceded financial crises, as Koetter and Poghosyan (2010) point out. They also argue that, due to decentralized trading with imperfect information and high transaction costs on the one hand and slow supply responses due to construction lags and limited land availability on the other, sustained deviations from the long-run equilibrium will occur more frequently in the housing market than in the financial markets.

In the housing market, the amount of credit made available by lenders depends on the net-worth of the debtors. Due to imperfections and informational asymmetries in the credit markets, a prospective borrower is usually granted a loan only by putting up collateral. In the models developed by Kiyotaki and Moore (1997) and Bernanke and Gertler (1989), shocks to the real economy are amplified through the credit market by altering the value of borrowers' net-worth.

This so-called *financial accelerator*[1] mechanism offers an explanation to the housing market fluctuations. First, higher housing prices increase the amount of credit needed to finance a given housing purchase. Thus, we would expect higher property valuations to put an upward pressure on the demand for credit. Second, most housing loans are secured by the property itself. An increase in housing prices raises the value of the housing capital, which feeds into a greater net-worth for the household sector. By increasing the net-worth and thus the value of the collateral, higher housing prices will increase their borrowing capacity. At the same time, higher property valuations make banks' assets less risky, as the increased value of the collateral pledged reduces the

[1]The term was coined in Bernanke and Gertler (1995), see also Bernanke et al. (1999).

likelihood of defaults on existing loans, which may motivate the banks to expand their lending.

That said, most housing purchases are financed by credit, and changes in household borrowing are expected to affect housing prices. The potential self-reinforcing mechanism that works between these markets makes it important to study from the perspective of financial stability, and it constitutes a main reason why central banks commonly assess financial sector vulnerability by monitoring both property prices and credit growth. The close relationship between the evolution of property prices and credit aggregates has been a focal point in the policy-oriented literature, see e.g. Borio et al. (1994).

In this paper, we analyze the interaction between housing prices and credit in Norway. The paper contributes to the literature in several ways. First, we use a system based cointegration analysis, while most existing studies rely on single-equation methods. We expect to find (at least) two cointegrating vectors and the system analysis is important for both identification and for estimation efficiency. The disposable income for the household sector is included as a third endogenous variable in the VAR and is found to be weakly exogenous with respect to the long-run coefficients in the model. This motivates why we focus on housing prices and credit in modeling the short-run adjustments.

Second, the dynamic interaction between housing prices and credit is also analyzed using system methods. Full information maximum likelihood is used in the design of the short-run specifications, which is carried out general-to-specific. Previous studies have resorted to an equation-by-equation approach at this stage.

Third, the paper includes a measure of households expectations about the future development in their own as well as the Norwegian economy in the dynamic specification. As a housing purchase is a long-term investment, this seems to be a highly relevant variable to include in a housing price equation. Indeed, it is shown that this variable has a positive and significant impact on housing prices.

While many previous studies have had difficulties measuring supply side effects, our results indicate a large and negative long-run impact on housing prices of an increase in the housing stock. This suggests that supply side constraints are important for long-run movements in prices and that a liberalization of zoning regulations and other regulations limiting the supply of housing might be an effective tool to prevent a rapid increase in housing prices.

Finally, dynamic simulations demonstrate how shocks are propagated and amplified across the two markets over time. When we take the analysis one step ahead and include a separate model for the supply side, the effects of a positive shock to housing prices or to credit are dampened over time as residential investments gradually shift the supply of housing.

The paper gives a survey of the recent literature in Section 2.2. A description of the Norwegian housing and credit markets is outlined in Section 2.3. Section 2.4 provides a brief theory discussion, while we investigate the fundamental determinants of housing prices and household debt in Section 2.5 by means of a system based cointegration analysis. Section 2.6 describes the dynamic interaction between the two variables. The model yields meaningful short and long-term effects when estimated on the sample 1986q2-2008q4. In Section 2.7, we compare our basic model for housing prices and household debt with an enlarged version which also includes the supply of housing. In both cases, dynamic simulations demonstrate that there are self-reinforcing feedback effects between the two variables of interest. Before concluding, Section 2.8 explores the robustness and stability of the model by adding four more years of data that have become available after the model was first documented.

2.2 A survey of empirical contributions

The empirical literature on housing prices is extensive; see e.g. Hendry (1984), Muellbauer and Murphy (1997), Pain and Westaway (1997), Meen (2001, 2002) and Malpezzi (1999) to mention a few important contributions. Girouard et al. (2006) provide a nice overview of the empirical literature. The majority of the papers have investigated the determinants of housing prices within a single-equation set-up. That framework does not shed light on the possible interaction between housing prices and household borrowing. Only recently – in the past decade – a literature on the nexus of housing prices and credit has emerged. The results up to now disagree about the direction of causality. The discrepancies can, however, be ascribed to a number of sources: there are institutional differences between countries, and the methodological approaches as well as sample sizes and data sets vary across the studies. A summary of the empirical findings on the interaction between housing prices (ph) and credit (d), which we refer to below, is given in Table 2.1 and Table 2.2.

In an early study, using both panel data and time series techniques for 20 countries, Hofmann (2003) finds a cointegrating relationship between property prices, bank lending and GDP. The equation is interpreted as a credit equation and property prices are found to affect private sector borrowing in the long-run, while the opposite direction of causation is not supported. The data are quarterly and cover the period 1985-2001. The author also reports results for the short-run dynamics, where he finds causality to go in both directions. The long-run results are further corroborated in Hofmann

Table 2.1: Literature Evidence on the Long-Run Interaction Between Housing Prices and Credit[a]

Author(s)	$ph \rightarrow d$	$ph \leftarrow d$	$ph \leftrightarrow d$
Hofmann (2003, 2004)	*		
Brissimis and Vlassopoulos (2009)	*		
Gerlach and Peng (2005)	*		
Oikarinen (2009a,b)		*	
Fitzpatrick and McQuinn (2007)			*
Berlinghieri (2010)			*
Gimeno and Martinez-Carrascal (2010)			*

[a] The table summarizes the literature evidence on the long-run interaction between housing prices and credit. Housing prices are denoted by ph, while credit is denoted by d.

Table 2.2: Literature Evidence on the Short-Run Interaction Between Housing Prices and Credit[a]

Author(s)	$ph \rightarrow d$	$ph \leftarrow d$	$ph \leftrightarrow d$
Hofmann (2003)			*
Brissimis and Vlassopoulos (2009)			*
Gerlach and Peng (2005)	*		
Oikarinen (2009a,b)[b]		*	
Fitzpatrick and McQuinn (2007)		*	
Berlinghieri (2010)			*

[a] The table summarizes the literature evidence on the short-run interaction between housing prices and credit. Housing prices are denoted by ph, while credit is denoted by d.

[b] The results apply to the period after the Finnish credit markets were deregulated.

(2004),[2] where he first studies VARs in real credit to the private sector, GDP (as a broad measure of economic activity) and the short-term real interest rate as a measure of financing costs for each country. For a majority of the countries, the Johansen analysis (Johansen (1988)) shows no cointegration with this information set. When he extends the analysis to include real property prices in the VARs, Hofmann finds strong support for one cointegrating vector for all countries, which (through the significance of the loadings) can be interpreted as a credit equation for those countries where a high share of loans are secured by real estate.

This finding is supported by Brissimis and Vlassopoulos (2009) in a single country study for Greece. With quarterly data specific to the housing market for the period 1993-2005, they find only one cointegrating relationship based on system based cointegration techniques. This is interpreted as a mortgage loan equation, where loans are determined by housing prices, interest rates and an income measure. The loadings reveal that only the credit variable equilibrium corrects, i.e. housing prices are found to be weakly exogenous with respect to the long-run parameters. Hence, in a long-run perspective,

[2] See also Goodhart and Hofmann (2007).

the causation does not run from mortgage lending to housing prices. In the short-run, they find evidence of a contemporaneous bi-directional dependence.

Gerlach and Peng (2005) examine the interaction between credit to the private sector and residential property prices with a sample of quarterly data for Hong Kong from 1984 to 2001. They use a vector equilibrium correction framework and find that the direction of causation is from housing prices to private sector debt both in the long-run and in the short-run.

Contrary to this, Oikarinen (2009b) finds the direction of causation to go from household borrowing to housing prices in the long-run. He uses quarterly data for Finland from 1975 to 2006 to explore the mutual dependence between housing prices and borrowing. A cointegration analysis in the spirit of Johansen (1988) supports the existence of only one cointegrating vector, which is interpreted as a housing price equation. Tests for Granger non-causality show that there is no dynamic effect going in either direction before 1988, i.e. before the Finnish credit market was considered fully deregulated. There is however an effect on housing prices from the credit market running via the equilibrium correction term. After the deregulation, however, lending is shown to Granger cause housing prices also through the short-run dynamics, while the opposite is not found to be the case. Furthermore, both variables are affected by the equilibrium correction term in the short-run after the deregulation has taken place. These results are corroborated by an impulse response analysis, where Oikarinen establishes an interaction between housing prices and credit only after the deregulation process was considered completed (after 1987). Using the same methodological framework, Oikarinen (2009a) reports similar results with regional housing price data for the Helsinki Metropolitan area. Again, household debt enters the long-run relationship for housing prices and Granger non-causality tests give the same results as in Oikarinen (2009b).

There are also a few recent studies documenting a mutual dependency in the long-run, i.e. two cointegrating vectors are found. Fitzpatrick and McQuinn (2007) look at the interaction between housing prices and mortgage credit in Ireland between 1981 and 1999. They show that the two variables are mutually dependent in the long-run, as well as in the short-run. In the dynamic specification, a contemporaneous effect is only established from credit to housing prices, while housing prices are found to have lagged effects on credit. Like Hofmann (2003), Fitzpatrick and McQuinn (2007) analyze the long-run dependence within a single-equation framework adopting the original approach to cointegration of Engle and Granger (1987).[3]

[3]Hofmann (2003) also considers a Johansen analysis, but it is the results from the single-equation procedure that are retained for the dynamic specifications.

When exploring the dynamic interaction between housing prices and credit, the two equations are estimated separately by OLS and a general-to-specific procedure is followed to find a parsimonious system. Acknowledging the potential endogeneity problems, Fitzpatrick and McQuinn estimate the two equations jointly by non-linear three stage least squares after having sequentially reduced the dimensionality of the two equations.[4]

The results of Fitzpatrick and McQuinn (2007) are supported by Berlinghieri (2010) for quarterly US data covering the period 1977 to 2005 who also finds a bi-directional interdependence in the long-run. A two step Engle-Granger approach is adopted and the short-run dynamics are estimated by single-equation OLS. The interaction is found to run in both directions also in the short-term.

Making use of quarterly data for the period 1984-2009, Gimeno and Martinez-Carrascal (2010) study the interaction between housing prices and household borrowing in Spain. A system based cointegration analysis shows that the two variables are interdependent in the long-run, i.e. housing prices affect mortgage credit in the long-run, and *vice versa*. Further, the loading factors imply that disequilibrium in the credit market leads to adjustments in both markets, while only housing prices equilibrium correct to disequilibrium constellations in the housing market. They do not report results for the short-run dynamics.

An alternative approach to modeling housing prices is adopted by Carrington and Madsen (2011), who consider a Tobin's Q model for US housing price determination over the sample 1967q2-2010q2. They use an ARDL bounds testing approach to test whether housing prices, the cost of agricultural land and construction costs are cointegrated. They do not find evidence for cointegration and consider a model in first differences instead. Interestingly, they find an important role of banks' willingness to lend for short-run fluctuations in housing prices. These results are confirmed by a panel analysis for eight OECD countries over the period 2003q1–2010q3.

The diverging results, as summarized in Table 2.1 and Table 2.2 call for further research. Our paper adopts the same econometric approach as Gimeno and Martinez-Carrascal (2010), but we go further. Not only do we to study the long-run interaction, but also the dynamic interaction between the two markets, which is important for both policy evaluation and forecasts.

The studies that address the short-run interaction by modeling the dynamics of the two variables all use a single-equation approach, i.e. the equations are estimated separately

[4] In addition to an equation for housing prices and one for household debt, Fitzpatrick and McQuinn (2007) adds an additional equation for the supply side of the housing market to their system. This equation is taken from a former study (McQuinn, 2004) and hence it is not directly integrated in their analysis.

by OLS regressions. In some cases, the system is estimated jointly by 3SLS after the dimensionality of the equations in the system have been reduced separately. This may be inappropriate – as pointed out by Hammersland and Jacobsen (2008) – because the single-equation specifications will themselves be affected by the reduction process if we believe the variables in the system are jointly determined in the first place. From this perspective, it seems highly relevant to deal with the potential simultaneity from the onset. Hence, one should design the structural short-run model using system methods that takes on the simultaneity problem from the outset.

2.3 The Norwegian housing and credit markets

The banking crisis in Norway that took place between 1988-1993 is a clear example of a collapse of property prices being followed by imbalances in the real economy. The recent financial crisis was different in that it was an external shock to the domestic economy, which had a significant, but short-lived, negative effect on Norwegian housing prices.

Krogh (2010) gives a detailed account of the changes in the Norwegian credit market regulations and other major events in the period 1970-2008. This time span entails a period with strict credit market regulations in the 1970s, a gradual deregulation of these markets in the 1980s, followed by the banking crisis, and the subsequent development up to the advent of the current financial crisis.

For our purpose, it is important to note that also the housing market was heavily regulated in Norway after World War II. Building materials were rationed and there were strict regulations on housing, both with regard to quantity and prices. These regulations ended in July 1982, with the abolition of price regulation on cooperative housing. The credit market regulations were lifted shortly after this. The combined effect of these liberalization processes was a boom in the real estate market, made possible and financed by a credit expansion. The problems facing the banking sector when the bubble burst became immense (Vale, 2004). After the Norwegian banking crisis, which ended in 1993, real housing prices have grown almost consecutively until the financial meltdown of the previous decade (see Figure 2.1a). Growing housing prices have been accompanied by a substantial expansion in real household debt (see Figure 2.1b).

The historical episodes referred to above strongly suggest there is an interdependency between the evolution of real housing prices and that of real household debt. For an impression of how housing price developments relate to the general macroeconomic picture in Norway, Figure 2.1c plots the four quarter growth in real housing prices against

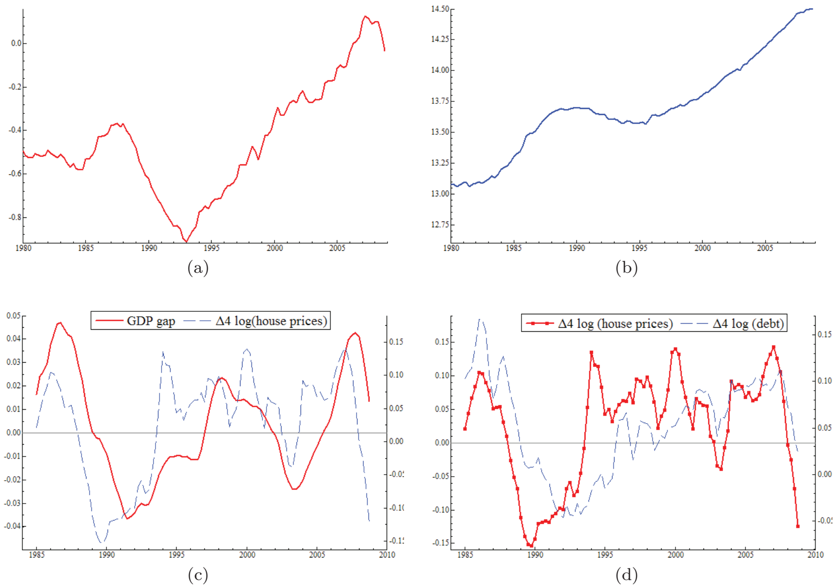

FIGURE 2.1: Panel a) Log of real housing prices, 1980-2008. Panel b) Log of real household debt, 1980-2008. Panel c) GDP gap (left scale) and four quarter growth in real housing prices (right scale), 1985-2008. Panel d) Four quarter growth in real housing prices (left scale) and in real household debt (right scale), 1985-2008. (*Source:* Statistics Norway)

percentage deviations of GDP mainland Norway from trend.[5] A close link between economic activity and housing prices is apparent over the entire period, with a less pronounced correlation pattern the last few years. Goodhart and Hofmann (2007) argue that there will be a tendency of changes in housing price growth to lead *peaks* and *troughs* in economic activity. This may suggest that turning points in the housing market are indicators of future economic developments. Figure 2.1c shows such a tendency for the case of Norway in the period after the deregulation of the Norwegian credit markets had been completed. Housing prices may affect economic activity through wealth effects on private consumption and a rise in house prices also raise the value of housing relative to construction costs, that is the Tobin q (Tobin, 1969) for residential investments. Another channel in which housing prices could have an effect on the business cycle is by amplifying shocks in the credit market. It is evident from Figure 2.1d, where we have plotted the four quarter growth in real housing prices against four quarter growth in real household borrowing, that the two series move quite closely together.

[5]GDP mainland Norway measures total production in Norway excluding two sectors: extraction of oil and gas, and ocean transport.

Previous studies of the credit and housing markets in Norway do not take the potential simultaneity between the two into account. For example, the determination of household debt is the topic of Jacobsen and Naug (2004), whilst Jacobsen and Naug (2005) describe a separate model for housing prices. In Jacobsen and Naug (2004), housing prices are one of the fundamental factors explaining household debt, whereas household borrowing is not part of the cointegrated vector explaining housing prices in Jacobsen and Naug (2005).[6] That said, it is documented that the interest rate is an important determinant of housing prices. Also, Jacobsen and Naug (2004) find that the interest rate is one of the fundamental factors explaining household borrowing. The effect of interest rates on credit thus suggests that the interest rate variable in the housing price equations captures a credit effect, i.e. the coefficient of the interest rate in Jacobsen and Naug (2005) picks up a gross effect.[7]

2.4 Economic theory

The commonly used framework for modeling housing prices is the life-cycle model, see e.g. Meen (2001, 2002), Muellbauer and Murphy (1997, 2008) and the references therein. We augment this model with a term capturing the presence of credit constraints, and the marginal rate of substitution (MRS) between housing and a composite consumption good is then given by (see e.g. Meen (1990) or Meen and Andrew (1998)):

$$MRS = PH_t \left[(1-\tau_t)i_t - \pi_t + \delta_t - \frac{\dot{PH}_t^e}{PH_t} + \lambda_t/\mu_c \right], \tag{2.1}$$

where PH_t is real housing prices, τ_t is the marginal tax rate on equity income, i_t is the nominal interest rate (paid by households for loans), π_t is the annual inflation rate, δ_t is the depreciation rate or the rate of maintenance costs including property taxation, and $\frac{\dot{PH}_t^e}{PH_t}$ is the expected real rate of appreciation for housing prices. λ_t is the shadow price of the credit constraint which is divided by the marginal utility of consumption μ_c. This is commonly known as the real housing user cost of capital, in this case augmented with a credit constraint. Market efficiency requires that the following no-arbitrage relationship

[6] Jacobsen and Naug (2005) tested for the significance of a credit variable in their specification, but found no significant effects.

[7] Akram et al. (2006), Akram et al. (2007) and Andersen (2011) augment the core part of a macroeconometric model for the Norwegian economy (see e.g. Bårdsen et al. (2003) and Bårdsen et al. (2005)) with different versions of the housing price and credit equations of Jacobsen and Naug (2004, 2005). These studies address issues related to financial stability when there are interaction effects between housing prices and credit.

holds, where Q_t represents the real imputed rental price for housing services

$$PH_t = \frac{Q_t}{(1-\tau_t)i_t - \pi_t + \delta_t - \frac{\dot{PH}_t^e}{PH_t} + \lambda_t/\mu_c} \tag{2.2}$$

Meen (2002) follows Poterba (1984) and interprets (2.2) as an inverted housing stock demand function. In the following, we will assume that the depreciation rate is constant.[8] If we assume that Q_t, which is unobservable, is a function of real disposable income for the household sector (excluding dividends), YH_t, and the stock of dwellings, H_t, we can write the inverted demand function as

$$PH_t = f^* \left(H_t, YH_t, R_t, \frac{\dot{PH}_t^e}{PH_t}, \lambda_t/\mu_c \right), \tag{2.3}$$

where $R_{t,}$ is the real after tax interest rate $(1-\tau_t)i_t - \pi$.

With a constant depreciation rate, the real user cost can be split in two different components: the real direct user cost (as measured by R_t) and expected real housing price appreciation. In the econometric analysis, we use the real direct user cost as our operational measure of the user cost and let price expectations be modeled by allowing lagged real price appreciation to enter our dynamic model.[9] This is similar to Abraham and Hendershott (1996), Gallin (2008) and Anundsen (2013) on US data, and it is consistent with the lagged housing price appreciation not having permanent effects, but rather that it picks up a momentum or the "bubble builder" effect using the terminology of Abraham and Hendershott (1996).[10]

Furthermore, we shall substitute household loans as a proxy for the theoretically correct – but unobservable – λ_t/μ_c term in (2.3).[11] Our empirical study can thus be seen as a test of the informational value of household loans when direct information on credit constraints is missing. As household debt is non-stationary, we implicitly assume that the same holds for the shadow price of the credit constraint.

[8]Assuming a constant depreciation rate is consistent with the Norwegian National accounts, where a constant depreciation rate is used for housing.

[9]It should be mentioned that we have experimented with a moving average process for the expectation component of the user cost. We find that this term is insignificant in our long-run relationships, suggesting that it is reasonable to assume that lagged price appreciation effects are picked up through the dynamics of the model. We then avoid making *a priori* assumptions about the expectation formation.

[10]Abraham and Hendershott (1996) distinguish between a bubble builder effect represented by lagged real housing price appreciation in the dynamic part of the model and a bubble burster effect through an equilibrium correction term.

[11]An alternative approach has been considered in Duca et al. (2011a,b) on US data. Including a measure of the LTV ratio for first time home buyers, they find that exogenous shifts in credit conditions have been important for US housing price dynamics in the 2000s.

Hence, we formulate the determination of real housing prices at the aggregate level in a static long-run equilibrium as

$$PH_t = f(H_t, YH_t, R_t, D_t), \tag{2.4}$$

where $\frac{\partial f}{\partial H} < 0, \frac{\partial f}{\partial YH} > 0, \frac{\partial f}{\partial R} \gtrless 0, \frac{\partial f}{\partial D} > 0$ and D_t is real household debt.

Equation (2.4) expresses market clearing prices for any given level of the housing stock. The equation describes housing prices as an increasing function of disposable income and household debt, while a greater supply of housing services is expected to push housing prices down. The sign of the derivative with respect to the interest rate is ambiguous. The main effects of a change in the interest rate work through disposable income and household loans, which both are controlled for in (2.4). What remains are the substitution effects which may be of either sign from a theoretical point of view.[12]

We supplement our model for housing prices with a relationship that determines real household debt in a long-run equilibrium

$$D_t = g(H_t, YH_t, R_t, PH_t, TH_t), \tag{2.5}$$

where $\frac{\partial g}{\partial H} > 0, \frac{\partial g}{\partial YH} > 0, \frac{\partial g}{\partial R} < 0, \frac{\partial g}{\partial PH} > 0, \frac{\partial g}{\partial TH} > 0$ and TH_t denotes the housing turnover. Equation (2.5) is an extended version of Fitzpatrick and McQuinn (2007). It defines household debt as a function of the housing stock, housing prices, the interest rate, disposable income and the housing turnover. In our specification, the housing stock and the housing turnover are additional explanatory variables. Since all the variables included in (2.4) and (2.5) are usually found to be non-stationary and integrated of first order, and since the theory postulates long-run equilibrium relationships, the discussion in this section suggests that housing prices and credit should be cointegrated with the variables – or a subset thereof – included in (2.4) and (2.5), i.e. we would expect to find two cointegrating relationships.

In the following we shall think of equations (2.4) and (2.5) as a subsystem, conditioning on H_t, YH_t, R_t, and TH_t. The last three variables can be assumed to be determined by factors other than housing prices and credit. The housing stock, H_t, on the other hand represents the supply side of the housing market. It appears in equation (2.3) since

[12] It is not only from a theoretical point of view that the sign of the direct effect is ambiguous. Empirically it is often found to be statistically insignificant. In the case of Norway the dominant interest rate effects on housing prices are indirect. Almost all mortgage debt in Norway are loans with flexible interest rates. Hence, a change in interest rates will immediately feed into the disposable income for households, and it is likely to pick up the main effect of interest rates on demand for housing. The inclusion of the credit aggregate captures the effect on housing prices from a change in the cost of financing.

it affects negatively the market clearing rent and hence the price of housing. We will assume it is related to the profitability of new construction and thus that it is influenced positively by real housing prices and negatively by construction costs. Hence, there are feedback effects from housing prices via H_t to housing prices and credit. In order to capture these feedback effects we estimate a submodel for housing supply separately in Appendix 2.A. In Section 2.7, when we compare the dynamic responses from our baseline model with those from an extended version of the model, which includes the housing supply, we find that the effects of a shock to housing prices or household debt are dampened.

2.5 Cointegration analysis

2.5.1 Methodological approach

A semi-logarithmic transformation of the variables appearing in equations (2.4) and (2.5) – which can be seen as a linearization of the theoretical formulations – forms the basis for the information set underlying our empirical analysis. All data are seasonally unadjusted and in what follows, small letters indicate that the variables are measured on a logarithmic scale.[13] All monetary variables are measured in real terms, having been deflated by the consumption deflator. Our sample covers the period 1986q2-2008q4. We have data for the number of housing transactions only from 1985q1, and the housing price data are also less reliable in the period prior to this. Since we consider a post-deregulation sample, it follows that we do not account for shifts in the constraints that are due to the deregulation of the Norwegian housing and credit markets. That said, the deregulation of the housing and credit markets in the early 1980's is likely to have altered the functioning of both, so that a different econometric model would probably be more suitable if we were to consider the period prior to the deregulation. In particular, it is less likely that a self-reinforcing relationship between housing prices and credit existed during the regulation period, since these regulations clearly distorted the ordinary market mechanisms.[14]

The orders of integration of the data series have been examined by a suite of different tests; the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller (1979)), the Phillips-Perron (PP) test (Phillips (1987) and Phillips and Perron (1988)), as well as

[13]For a detailed data description, see Appendix 2.B. The log transformation is applied to all variables in (2.4) and (2.5), except the real after tax interest rate.

[14]This is consistent with the empirical findings of Oikarinen (2009b), who finds that a two-way interaction between housing prices and credit in Finland can only be established after liberalization of the credit markets in the late 1980s.

the Kwiatowski, Phillips, Schmidt, and Shin (KPSS) test (Kwiatkowski et al. (1992)).[15] Based on these tests, we treat all variables as integrated of order one at most in the econometric analysis. There is also supporting evidence for this approach in that we find - as we report below - that the residuals in the final empirical model turn out to be stationary. Details on the tests for unit roots are given in Table 2.C.1 of Appendix 2.C.

Due to the non-stationarity of the variables in our data set, we start by investigating the the long-run determinants of housing prices and household borrowing in a cointegrated VARX system where also household income is treated as an endogenous variable, while we condition on the real after tax interest rate, the housing turnover and the housing stock. Finding evidence of cointegration ensures that we can formulate the VARX as a vector equilibrium correction model (VECM). The VECM approach provides an opportunity to study long-run determinants and short-run dynamics in a unified framework, which opens for the possibility that the causality between housing prices and credit is bi-directional both in the short-run and in the long-run. The model is therefore suitable for addressing the key issue: is there empirical evidence for the existence of a financial accelerator in the Norwegian housing market?

In general, the I(1) cointegrated VAR (CVAR) model can be written as a re-parameterization of a $VAR(p)$ model, see for example Johansen (1988), Johansen (1995) and Juselius (2006):

$$\Delta \boldsymbol{Y}_t = \boldsymbol{\Pi} \boldsymbol{Y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \boldsymbol{Y}_{t-1} + \boldsymbol{\Phi} \boldsymbol{D}_t + \boldsymbol{\varepsilon}_t, t = 1, ..., T \tag{2.6}$$

$\boldsymbol{Y}_t$ is a $n \times 1$ matrix comprising the endogenous variables in the system, while $\boldsymbol{D}_t$ contains deterministic terms such as a constant, linear trends or other regressors considered to be fixed. We let $\boldsymbol{\Pi}$, $\boldsymbol{\Gamma}_i$ and $\boldsymbol{\Phi}$ denote the coefficient matrices. With reference to a $VAR(p)$ model, the $\boldsymbol{\Pi}$ and $\boldsymbol{\Gamma}_i$ matrices are defined as $\boldsymbol{\Pi} = \sum_{i=1}^{p} \boldsymbol{\Pi}_i - \boldsymbol{I}$ and $\boldsymbol{\Gamma}_i = -\sum_{j=i+1}^{p} \boldsymbol{\Pi}_j$, where $\boldsymbol{\Pi}_i$ is the VAR coefficient matrix attached to lag number i. The innovation terms, $\boldsymbol{\varepsilon}_t$, are assumed to be independently Gaussian distributed, $N(\boldsymbol{0}, \boldsymbol{\Sigma})$, and the initial values $Y_{-p}, ..., Y_0$ are considered fixed.

In our case, we consider a $VARX(p, q)$, i.e. some of the variables in the system are treated as weakly exogenous. In addition, we follow the suggestion of Harbo et al. (1998) for partial systems and restrict a deterministic trend to enter the cointegration

[15] As a guidance for choosing the optimal lag truncation for the ADF test, we have relied on Akaike's information criterion (AIC) starting with an initial lag length of eight in the first differences in all test regressions and then chosen the specification with the lowest AIC value.

space. Thus, the $VECM(p,q)$ representation of the $VARX(p,q)$ that forms the basis for our econometric analysis reads:

$$\Delta \boldsymbol{X}_t = \tilde{\boldsymbol{\Pi}}\tilde{\boldsymbol{Y}}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \boldsymbol{X}_{t-i} + \sum_{i=0}^{q-1} \boldsymbol{\Psi}_i \Delta \boldsymbol{Z}_{t-i} + \tilde{\boldsymbol{\Phi}}\tilde{\boldsymbol{D}}_t + \boldsymbol{\varepsilon}_t. \tag{2.7}$$

where $\boldsymbol{X}_t$ is a 3×1 matrix comprising the endogenous variables ph, d and y, while $\boldsymbol{Y}_t = (\boldsymbol{X}_t', \boldsymbol{Z}_t')'$ is a $(3+3) \times 1$ matrix where $\boldsymbol{Z}_t$ is a 3×1 matrix composed of the weakly exogenous variables R, th and h and $\tilde{\boldsymbol{Y}}_t = (\boldsymbol{Y}_t', t)'$ with t denoting a deterministic trend. The vector $\tilde{\boldsymbol{D}}_t$ comprise a constant and centered seasonal dummies.

The trace test for the order of cointegration (Johansen, 1988) can be used to determine the rank of the matrix $\tilde{\boldsymbol{\Pi}}$, which corresponds to the number of independent linear combinations between the variables that are stationary. We follow Johansen (1988) and define $\tilde{\boldsymbol{\Pi}} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\beta}$ is a $(n+k+1) \times r$ matrix and $\boldsymbol{\alpha}$ is a $n \times r$ matrix corresponding to the long-run coefficients and loading factors respectively. The rank of the $\tilde{\boldsymbol{\Pi}}$ matrix is denoted by r, while n refers to number of endogenous variables and $k+1$ is the number of exogenous variables (including the deterministic trend, which is restricted to lie in the cointegration space). Thus, in our case – with $n = k = 3$ – $\boldsymbol{\beta}$ is a $7 \times r$ matrix and $\boldsymbol{\alpha}$ is a $3 \times r$ matrix.

2.5.2 Cointegration results

As mentioned, our starting point for the cointegration analysis is a VARX in real housing prices, real household debt and real disposable income, while we condition on the real after tax interest rate, the housing turnover and the housing stock.[16] We start with a lag length of 5 in both the endogenous and the weakly exogenous variables ($p = q = 5$), which ensures that we have a well specified model without evidence of autocorrelation, heteroskedasticity nor non-normality. Then, the optimal lag truncation is decided based on AIC. According to AIC, the VAR-model should include 5 lags in the endogenous variables, while we find that only one lag is needed for the weakly exogenous variables.[17]

[16]Indeed, including the turnover as an endogenous variable in the VAR, we find that it is weakly exogenous (the p-value from the test is 0.6847). This supports our conditioning and saves valuable degrees of freedom. Alternatively, weak exogeneity can be tested along the lines of Johansen (1992), Harbo et al. (1998), Pesaran et al. (2004) and Dees et al. (2007), i.e. by including the two cointegrating vectors we document below in the marginal model for the turnover and then test their joint significance. An F-test of the two zero restrictions has a p-value of 0.1891, which gives further justification to this assumption.

[17]Details are available in Table 2.C.2 in Appendix 2.C.

Having decided on the lag length, we use the trace test to decide on the number of cointegrating relationships. Table 2.3 displays the results. We find that there are two cointegrating vectors.[18] The model is well specified – residual diagnostics show that the residuals are neither heteroskedastic nor autocorrelated, and normality is not rejected.

TABLE 2.3: Trace test for cointegration [a]

$Eigenvalue : \lambda_i$	H_0	H_A	λ_{trace}	5%-critical value[b]
0.39	$r = 0$	$r \geq 1$	86.59	64.48
0.22	$r \leq 1$	$r \geq 2$	41.74	40.95
0.19	$r \leq 2$	$r = 3$	18.82	20.89
Diagnostics[c]	Test statistic	Value[p-value]		
Vector AR 1-5 test:	F(45,146)	1.06 [0.39]		
Vector Normality test:	$\chi^2(6)$	7.78 [0.26]		
Vector Hetero test:	F(270,247)	1.03 [0.42]		
Estimation period:	1986q2-2008q4			

[a] Endogenous variables: Real housing prices (ph), real household debt (d) and real disposable income (yh). Restricted variables: Real interest rate after tax (R), housing turnover (th), housing stock (h) and a trend (t). Unrestricted variables: Constant and centered seasonal dummies for the first three quarters.
[b] Critical values are obtained from Table 13 in Doornik (2003) - with 3 exogenous variables.
[c] See Doornik and Hendry (2009a).

Exact identification can be achieved by imposing two restrictions in each vector. We start by normalizing on real housing prices in the first vector and real household debt in the other. In addition, it is assumed that the housing turnover has no direct effect on real housing prices.[19] This is in accordance with the theoretical housing price equation (2.4), while earlier studies have found that the turnover affects household borrowing in Norway (see Jacobsen and Naug (2004)), which suggests that it should be part of the relationship determining household debt. The final restriction we use for exact identification is that it is the value of the housing capital – and not simply housing prices – which determines the size of the collateral. To incorporate this into the empirical framework, we assume that a change in either the housing stock or housing prices have the same effect on household debt.

Based on the identified cointegrated vectors, we can move on to test overidentifying restrictions. The results for these restrictions are documented in Table 2.4 below.[20] For every new restriction that is imposed, we report both the log-likelihood value, the incremental test as well as the total test at the bottom line of each panel. In Panel 1, the trend variable is dropped from both equations, which correspond to two testable overidentifying restrictions. Next, in Panel 2, we omit the real after tax interest rate

[18] Critical values correcting for the inclusion of exogenous variables (see Doornik (2003)) have been used.

[19] Gimeno and Martinez-Carrascal (2010) and Fitzpatrick and McQuinn (2007) exclude the real interest rate from the long-run equation for housing prices by assumption. Pursuing this alternative identification strategy, i.e. excluding the real interest rate instead of the turnover from the housing price equation from the outset, we get identical results to those reported below.

[20] The absolute value of standard errors are reported in parentheses below the estimated coefficients.

from the vector associated with real housing prices. As mentioned above, this does not imply that a change in the interest rate will not affect housing prices, but it means that interest rate effects are captured by changes in disposable income and through the credit channel. In Panel 3, there is no effect of disequilibrium in the housing market on household debt, whereas Panel 4 shows the case with no direct effect of real disposable income on household debt. Finally, Panel 5 shows the result when we impose that the loadings of both cointegrating vectors with respect to income are zero, i.e. the test shows weak exogeneity of income with respect to the long-run coefficients, see Johansen (1992). According to the incremental tests reported in Table 2.4, all individual restrictions are supported by the data and the p-value for the joint test of all restrictions is 0.3.

TABLE 2.4: Testing steady-state hypotheses.

The just identified house price and debt equations are defined by
$ph = \beta_{d,1}d + \beta_{yh,1}yh + \beta_{h,1}h + \beta_{R,1}R + \beta_{t,1}t$
$d = \beta_{ph,2}ph + \beta_{yh,2}yh + \beta_{R,2}R + \beta_{th,2}th + \beta_{h,2}h + \beta_{t,2}t$
Panel 1: Testing no trend ($\beta_{t,1} = \beta_{t,2} = 0$)
$ph = \underset{(0.07)}{0.76}\,d + \underset{(0.21)}{1.39}\,yh - \underset{(0.37)}{2.00}\,h + \underset{(0.85)}{0.13}\,R$
$d = 1.53ph - \underset{(0.17)}{1.45}\,yh - \underset{(1.40)}{0.71}\,R + \underset{(0.05)}{0.09}\,th + \underset{(0.07)}{1.53}\,h$
$LogL = 842.845$, $\chi^2(2) = 3.81[0.15]$
Panel 2: No effect of real after tax interest rate on house prices ($\beta_{R,1} = 0$)
$ph = \underset{(0.08)}{0.77}\,d + \underset{(0.22)}{1.43}\,yh - \underset{(0.40)}{2.07}\,h$
$d = 1.54ph - \underset{(0.18)}{1.48}\,yh - \underset{(0.40)}{0.54}\,R + \underset{(0.05)}{0.10}\,th + \underset{(0.07)}{1.54}\,h$
$LogL = 842.834$, $\chi^2(1) = 0.02[0.88]$, $\chi^2(3) = 3.84[0.28]$
Panel 3: No effect of disequilibrium housing prices on household debt
$ph = \underset{(0.19)}{0.84}\,d + \underset{(0.65)}{1.67}\,yh - \underset{(1.18)}{2.58}\,h$
$d = 1.08ph - \underset{(0.85)}{1.18}\,yh - \underset{(2.35)}{3.98}\,R + \underset{(0.28)}{0.56}\,th + \underset{(0.30)}{1.08}\,h$
$LogL = 842.276$, $\chi^2(1) = 1.12[0.29]$, $\chi^2(4) = 4.95[0.29]$
Panel 4: No effect of real disposable income on household debt ($\beta_{yh,2} = 0$)
$ph = \underset{(0.19)}{0.86}\,d + \underset{(0.64)}{1.42}\,yh - \underset{(1.16)}{2.33}\,h$
$d = 0.78ph - \underset{(1.87)}{2.83}\,R + \underset{(0.15)}{0.24}\,th + \underset{(0.15)}{0.78}\,h$
$\mathrm{LogL} = 841.323$, $\chi^2(1) = 1.12[0.29]$, $\chi^2(5) = 6.86[0.23]$
Panel 5: Imposing weak exogeneity of income with respect to the long-run coefficients :
$ph = \underset{(0.19)}{0.98}\,d + \underset{(0.63)}{1.69}\,yh - \underset{(1.15)}{3.03}\,h$
$d = 0.76ph - \underset{(1.79)}{2.74}\,R + \underset{(0.15)}{0.28}\,th + \underset{(0.16)}{0.76}\,h$
$\alpha_{1,ph} = \underset{(0.04)}{-0.24}, \alpha_{1,d} = \underset{(0.03)}{-0.10}, \alpha_{2,d} = \underset{(0.01)}{-0.04}$
$LogL = 840.529$, $\chi^2(2) = 1.59[0.451]$, $\chi^2(7) = 8.44[0.30]$
The sample is 1986q2 to 2008q4, 91 observations.

Note: For notation, confer footnote a in Table 2.3 and the variable definitions in Appendix 2.B.

The coefficients reported in Panel 5 in Table 2.4, describe the two final long-run relationships for housing prices and household debt.[21] Our results support the hypothesis

[21] In Table 2.C.3 in Appendix 2.C, we report the loading factors corresponding to each of the panels.

that housing prices and household borrowing are mutually dependent in the long-run. All long-run coefficients have the expected signs in the final model (Panel 5) and they are significant at conventional significance levels.[22]

The semi-elasticity of household borrowing with respect to the real interest rate after tax is -2.74, implying that a one percentage point increase in the real interest rate will decrease household borrowing by almost three percent in the long-run. This is lower (in absolute value) than the estimate found for Spain by Gimeno and Martinez-Carrascal (2010) who consider nominal instead of real interest rates. It is however greater than the estimates found by Brissimis and Vlassopoulos (2009) for Greece and Fitzpatrick and McQuinn (2007) for Ireland who both consider real interest rates. Even though there is no direct causal link between real housing prices and the real interest rate in our model, a higher interest rate implies that housing prices will fall as it reduces the demand for housing by altering the credit variable, which is found to be highly significant in the housing price equation.

The estimated elasticity of housing prices with respect to household debt is 0.98. This is lower than the elasticity reported by Fitzpatrick and McQuinn (2007), but higher than the estimate in Gimeno and Martinez-Carrascal (2010). We find that the credit aggregate exercises a greater impact on housing prices than do housing prices on credit in a long-run perspective, a result that parallels the finding of Fitzpatrick and McQuinn (2007). A one percent increase in housing prices will increase household borrowing by 0.76 *percent* in the long-run.

The adjustment coefficients (confer Panel 5) imply that both housing prices and household debt equilibrium correct when the latter departs from the value implied by its fundamentals ($\alpha_{1,d} = -0.1$ and $\alpha_{2,d} = -0.04$). Moreover, the analysis indicates that only housing prices equilibrium correct when housing prices are deviating from their steady state level ($\alpha_{1,ph} = -0.24$). This result is supported by Gimeno and Martinez-Carrascal (2010) for the case of Spain. It is interesting to note that housing prices are adjusting more rapidly to equilibrium than household debt. This is because the volume of debt is not that easily changed over night.[23]

It is worth emphasizing that our results does not suggest any separate population effects on neither housing prices nor household borrowing. This can easily be seen by

[22]The interest rate is the only exception. However, using a one sided test, which appears to be meaningful, it is found to be significant at the 10% level (p-value = 0.068). The fact that it is also highly significant from an economic point of view suggests that it should not be excluded.

[23]While we have only reported the adjustment coefficients from the final long-run relationships in Table 2.4, Table 2.C.3 in Appendix 2.C reports the adjustment coefficients corresponding to all the panels in Table 2.4.

reparameterizing the two cointegrating relationships in per capita terms.

$$ph = \beta_{d,1}\frac{d}{pop} + \beta_{yh,1}\frac{yh}{pop} + \beta_{h,1}\frac{h}{pop} + (\beta_{d,1} + \beta_{yh,1} + \beta_{h,1})\, pop$$
$$\frac{d}{pop} = \beta_{ph,2}ph + \beta_{R,2}R + \beta_{th,2}th + \beta_{h,2}\frac{h}{pop} + (\beta_{h,2} - 1)\, pop$$

where *pop* is log population. Thus, for the model to imply no additional population effects, the two additional restrictions that $\beta_{d,1} + \beta_{yh,1} + \beta_{h,1} = 0$ and $\beta_{ph,2} = \beta_{h,2} = 1$ need to hold. Imposing these two restrictions gives a p-value of 0.2449 for all nine restrictions imposed on the system, while the partial test for the two restrictions has a p-value of 0.2203. Thus, we can conclude that there is no loss of generality from not including a separate population variable in the model, which save us valuable degrees of freedom.

To investigate the recursive stability of the two long-run relationships, we have estimated the model quarter-by-quarter over the period 2000q1–2008q4. The recursively estimated coefficients are shown in Figure 2.2. It is clear that all the long-run coefficients in both vectors are fairly stable when estimated recursively. The lower left panel shows the recursively estimated likelihood ratio statistic[24] against the 5% critical value from the $\chi^2(7)$ distribution, and we see that the restrictions are accepted recursively as well.

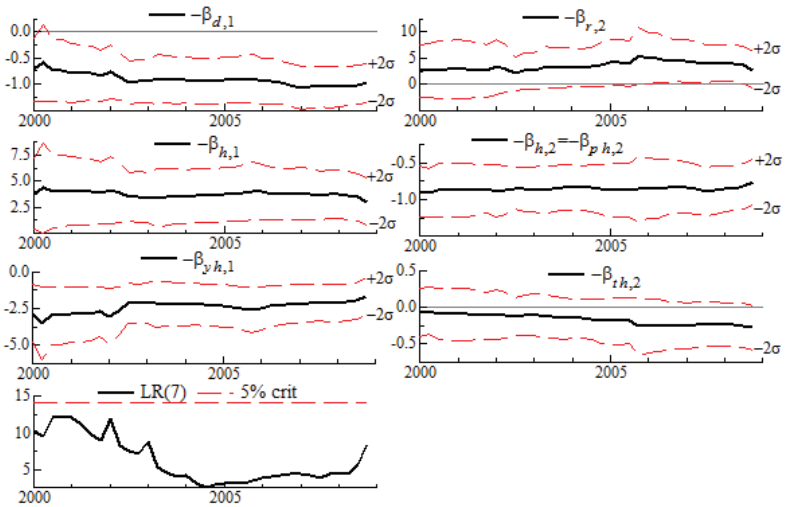

FIGURE 2.2: Recursively estimated coefficients and likelihood ratio test, 2000q1–2008q4

[24]The unrestricted likelihood ($LogL_{UR}$) is derived from the model in Panel 1 of Table 2.4, while the restricted likelihood ($LogL_R$) is based on the model reported in Panel 5. The likelihood ratio statistic is then calculated as $-2\,(LogL_R - LogL_{UR})$.

2.6 Short-run dynamics

2.6.1 Methodological approach

To derive the simultaneous equation system, the structural vector equilibrium correction model (SVECM), that forms the basis for the analysis of the short-run dynamics, we premultiply the reduced form representation in (2.7) by the (non-zero) contemporaneous feedback matrix, $\boldsymbol{B}$:

$$\boldsymbol{B}\Delta\boldsymbol{X}_t = \boldsymbol{B}\tilde{\boldsymbol{\Pi}}\tilde{\boldsymbol{Y}}_{t-1} + \sum_{i=1}^{4}\boldsymbol{B}\boldsymbol{\Gamma}_i\Delta\boldsymbol{X}_{t-i} + \sum_{i=0}^{4}\boldsymbol{B}\boldsymbol{\Psi}_i\Delta\boldsymbol{Z}_{t-i} + \boldsymbol{B}\boldsymbol{\Phi}\boldsymbol{D}_t + \boldsymbol{B}\boldsymbol{\epsilon}_t \tag{2.8}$$

where we now define $\boldsymbol{B}\tilde{\boldsymbol{\Pi}} = \boldsymbol{B}\boldsymbol{\alpha}\boldsymbol{\beta}' = \boldsymbol{\alpha}^*\boldsymbol{\beta}'$, $\boldsymbol{B}\boldsymbol{\Gamma}_i = \boldsymbol{\Gamma}_i^*$, $\boldsymbol{B}\boldsymbol{\Psi}_i = \boldsymbol{\Psi}_i^*$, $\boldsymbol{B}\boldsymbol{\Phi} = \boldsymbol{\Phi}^*$, $\boldsymbol{B}\boldsymbol{\epsilon}_t = \boldsymbol{\varepsilon}_t$. The new error term will also be IIN with zero mean and variance-covariance matrix given by: $\boldsymbol{\Omega} = E(\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t') = \boldsymbol{B}E(\boldsymbol{\epsilon}_t\boldsymbol{\epsilon}_t')\boldsymbol{B}' = \boldsymbol{B}\boldsymbol{\Sigma}\boldsymbol{B}'$.

As the income variable was found to be weakly exogenous, we can write the above system as a conditional system for housing prices and credit and a marginal model for income (see e.g Johansen (1992)). Since the focus of our paper is the interaction between housing prices and credit, we can, without loss of generality, abstract from modeling the marginal model for income. In that case, the conditional SVECM takes the following form:

$$\begin{aligned}
\Delta ph_t - b_{12}\Delta d_t = & \sum_{i=1}^{4}\boldsymbol{\Gamma}_{1i}^*\Delta\boldsymbol{X}_{t-i}^* + \sum_{i=0}^{4}\boldsymbol{\Psi}_{1i}^*\Delta\boldsymbol{Z}_{t-i}^* + \sum_{i=1}^{4}\tilde{\Psi}_{1,R_i}\Delta R_{t-i} \qquad (2.9)\\
& + \boldsymbol{\Phi}_1^*\boldsymbol{D}_t + \alpha_{1,ph}^* ECM_{t-1}^{ph} + \alpha_{1,d}^* ECM_{t-1}^{d} + \varepsilon_{ph,t}\\
-b_{21}\Delta ph_t + \Delta d_t = & \sum_{i=1}^{4}\boldsymbol{\Gamma}_{2i}^*\Delta\boldsymbol{X}_{t-i}^* + \sum_{i=0}^{4}\boldsymbol{\Psi}_{2i}^*\Delta\boldsymbol{X}_{t-i}^* + \sum_{i=1}^{4}\tilde{\Psi}_{2,R_i}\Delta R_{t-i} \qquad (2.10)\\
& + \boldsymbol{\Phi}_2^*\boldsymbol{D}_t + \alpha_{2,ph}^* ECM_{t-1}^{ph} + \alpha_{2,d}^* ECM_{t-1}^{d} + \varepsilon_{d,t}
\end{aligned}$$

where we have normalized such that the contemporaneous feedback matrix, $\boldsymbol{B}$, has ones along the main diagonal. $\boldsymbol{X}_t^*$ now consists of the two remaining endogenous variables, while $\boldsymbol{Z}_t^*$ still represents a vector of the weakly exogenous variables in the system (including the income variable). The constant and the centered seasonal dummies are collected in $\boldsymbol{D}_t$. $\boldsymbol{\Gamma}_{ji}^*$, $\boldsymbol{\Psi}_{ji}^*$, $\tilde{\Psi}_{j,R_i}$ and $\boldsymbol{\Phi}_j^*$ (j=1, 2) are the short-run coefficients, where $\boldsymbol{\Gamma}_i^* = (\boldsymbol{\Gamma}_{1i}^*, \boldsymbol{\Gamma}_{2i}^*)$,$\boldsymbol{\Psi}_i^* = (\boldsymbol{\Psi}_{1i}^*, \boldsymbol{\Psi}_{2i}^*)$ and $\boldsymbol{\Phi}^* = (\boldsymbol{\Phi}_1^*, \boldsymbol{\Phi}_2^*)$. Since the housing stock adjusts slowly, it is assumed to be fixed in the short-run and is not part of the vector $\boldsymbol{Z}_t^*$. Note also that we have excluded the contemporaneous value of the change in real after-tax interest rate, ΔR_t, from both equations to form our general unrestricted model. However, we supplement the short-run dynamics by including an expectations variable, E, which measures households expectations about future developments in their personal economy and the macroeconomy. Hence, $\boldsymbol{Z}_t^* = (th, E, yh)$. This is the system that constitutes the general unrestricted model. This variable can also be considered as a proxy for the expected rate of appreciation in housing prices, cf. Section 2.4.[25]

2.6.2 Results for dynamic model

The simultaneous equation system represented by (2.9) and (2.10) is estimated and designed simultaneously, and once again we have to face the tough and non-trivial decision of how to exactly identify the system. To achieve exact identification, we have chosen to exclude the contemporaneous effect of the turnover in the housing price equation, while the credit equation is identified by omitting the contemporaneous value of the expectations variable. The just identified system is estimated by FIML (full information maximum likelihood). The resulting model produces well behaved residuals and serves as a starting point for the reduction process to obtain a parsimonious representation of the system.

A parsimonious model is found by stepwise elimination of insignificant variables in the system, which are excluded either one by one or in blocks. Unlike the single-equation case, no algorithm for automatic general-to-specific search exists as yet, so we have carried out the search manually.[26] In that process, we make sure that, according to the diagnostic tests, the Gaussian properties of the residuals are retained and that all imposed restrictions are supported by the data. In the preferred (final) model, we have chosen to retain the income variable in the credit equation, which is relevant from *a priori* theoretical considerations, although it should have been excluded at the early stages of the reduction process had we followed a strict general-to-specific procedure. By doing so, we have achieved a more theoretically and intuitively appealing model formulation than we would have obtained otherwise, i.e if we had systematically eliminated the most insignificant variable at each stage. This procedure of structural model design results in the specifications displayed in Table 2.5.[27]

[25] The expectations variable is only available from 1992q3 and is set to 0 in the period prior to this. The expectations variable has previously been adopted by Jacobsen and Naug (2005). They find a positive and significant short-run effect of expectations on housing prices in a single-equation framework.

[26] See Doornik (2009) for a description of the automatic specification search in the case of a single-equation.

[27] Unlike previous studies (cf. Fitzpatrick and McQuinn (2007) and Brissimis and Vlassopoulos (2009)), the top-down approach applied in this paper consists of modeling the system simultaneously at all steps in the reduction process. Another approach, commonly used in the literature, is instead to simplify the two equations individually before estimating them as a system. Comparing our results to the results

Table 2.5: Short-run dynamics [a]

	Real housing prices		Real household debt	
Variable	Coefficient	t-value	Coefficient	t-value
Constant	1.542	7.71	0.048	6.39
Δd_t	0.859	2.25	-	-
Δd_{t-1}	-	-	0.173	1.88
Δd_{t-3}	0.309	2.32	-	-
Δph_{t-4}	0.389	4.88	-	-
Δyh_{t-3}	-	-	0.197	3.31
ΔE_t	0.093	4.40	-	-
ΔE_{t-1}	0.098	4.41	-	-
ΔE_{t-2}	0.055	2.40	-	-
ΔR_{t-4}	-	-	-0.258	2.16
ECM^{ph}_{t-1}	-0.175	7.82	-	-
ECM^{d}_{t-1}	-0.059	2.23	-0.046	6.11
Dummy, q1	0.022	3.75	-0.004	1.18
Dummy, q2	0.021	3.65	-0.00001	0.02
Dummy, q3	0.012	2.05	-0.007	2.05
Sargan		$\chi^2(46) =$	55.79 [0.1528]	
Log likelihood		560.26		
σ	0.0143		0.0098	
Diagnostics[b]		Test statistic	Value [p-value]	
Vector EGE-AR 1-5 test:		F(20,140)	0.90 [0.59]	
Vector Normality test:		$\chi^2(4)$	5.34 [0.25]	
Vector hetero test:		F(183,81)	0.88 [0.76]	
Estimation Method	FIML			
Sample	1986q2-2008q4 ($T = 91$)			

[a] Absolute t-values are reported.
[b] See Doornik and Hendry (2009a).

Table 2.5 reveals that credit effects are important for housing price fluctuations also in the short-run. We do not find any direct short-run effect running from housing prices to household debt though. It is however clear that the credit aggregate will be influenced by housing prices through the equilibrium correction term present in the credit equation. This means that it takes about one quarter before a shock to housing prices is transmitted to the credit market. Consistent with the cointegration analysis, the short-run analysis indicates that both housing prices and household debt equilibrium correct when household debt is high relative to its stable long-run equilibrium and that only housing prices equilibrium correct when departing from their fundamentals. Our results suggest that if housing prices depart from their long-run equilibrium by one percent, housing prices will fall by -0.175 *percent*. This is greater than what is found by Jacobsen and Naug (2005),[28] but lower than the estimate reported by Fitzpatrick and McQuinn (2007).

we would have obtained following this approach, we find that the methodology followed in this paper produces results that are both more reasonable and easier to interpret from an economic point of view. Details are available in Appendix 2.D.

[28] Jacobsen and Naug (2005) only consider housing prices and not the interaction between housing prices and household debt.

Like Jacobsen and Naug (2004, 2005) and Fitzpatrick and McQuinn (2007), we find that the credit aggregate has a slower adjustment towards equilibrium when it is departing from its fundamentals than do housing prices. This is not a very surprising finding in light of the fact that the volume of debt is not easily changed over night. Gimeno and Martinez-Carrascal (2010), however, find the opposite to be the case for Spain.

All estimated coefficients have the expected signs. Interestingly, we find that changes in expectations have a great impact on housing prices. The full effect is reached after three quarters, i.e., when there has been a change of 'mood'. As anticipated, our estimation results show that the interest rate has a negative impact on household borrowing (and therefore indirectly on housing prices) and the income variable lagged three quarters enters the credit equation significantly with an expected positive sign. As the equilibrium correction term for household debt is present in the housing price equation, the interest rate feeds into housing prices also here. The diagnostics indicate that the model is well specified and we find support for the imposed restrictions (p-value = 0.1528). The residuals from the two estimated equations are clearly stationary (see Table 2.C.4 in Appendix 2.C.

In Figure 2.3, we have plotted *ex ante* dynamic forecasts for the two endogenous variables. The forecasts are conditional on the explanatory variables as they accrued. The model does not fare too bad in *ex ante* forecasting, with two exceptions: the model under predicts the rapid recovery of house prices in the first half of 2009. More importantly, the credit forecasts are outside the forecast confidence bands in 2010q1 and 2011q1. However, this can for a large part be attributed to extremely cold winters, which lead to an extraordinary jump in electricity prices in each of those quarters and thus affected the consumption deflator we have used for the nominal-to-real transformations.

To explore this formally, Figure 2.4 shows *ex ante* dynamic forecasts for the two variables based on a slightly modified version of the model, where we have de-restricted the short-run price homogeneity.[29] Hence, we included the change in the price deflator, contemporaneously and at the first lag, in the short-run model. While both could be excluded from the housing price equation, both were significant with opposite signs in the credit equation. In fact, we can not reject the hypothesis that the two coefficients are equal in absolute value, i.e. suggesting that this captures a surprise inflation. As is seen, the forecasting accuracy of the model is improved. The forecasts for credit growth in 2010q1 and 2011q1 are no longer outside their confidence bounds.

[29]Details on the alternative forecasting model are available in Appendix 2.E.

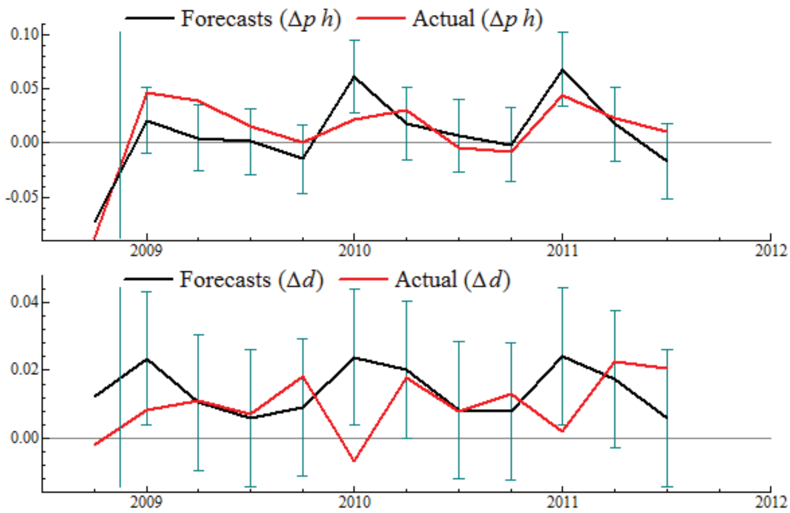

FIGURE 2.3: Ex ante forecasts from the "baseline" model, 2009q1–2011q3

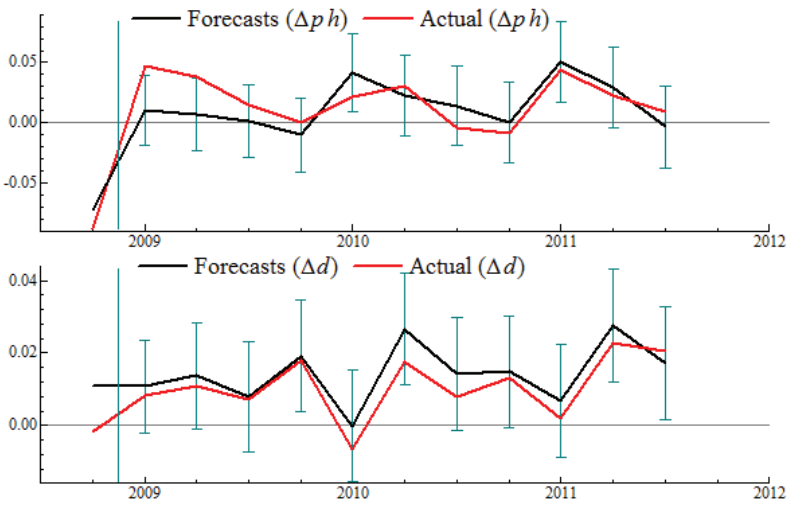

FIGURE 2.4: Ex ante forecasts from the model *without* short-run price homogeneity, 2009q1–2011q3

2.7 Dynamic effects of shocks

In the previous section, we used a general-to-specific approach to specify a parsimonious system capturing the dynamic interaction between housing prices and credit. In the following we will use Monte Carlo simulations of this model to show the dynamic responses to exogenous shocks to the system. As a first step, we consider the subsystem of housing prices and credit developed in Section 2.7.1, where we condition on the supply side of the housing market. In Section 2.7.2, we augment the subsystem with a small model for the supply side of the Norwegian housing market. This model is simply taken from an existing model for the Norwegian economy, i.e. the Statistics Norway forecasting model KVARTS, see Appendix 2.A for details. As will become evident in the following subsections, including the supply side dampens the long-run impact of shocks, as construction activity responds to changes in housing prices.

2.7.1 Dynamic multipliers: The baseline model

The first set of simulations we perform are based on the subsystem of housing prices and credit presented in Section 2.6. All simulations are conducted using 1000 stochastic Monte Carlo replications and 95 percent simulated confidence intervals (dotted red lines) are reported along with the simulated response path (solid blue lines). The dynamic effects of a permanent increase in the growth of credit and housing prices are shown in Figure 2.5 and 2.6, respectively. The figures display the impact on the growth rates as well as on the level of real housing prices and the stock of real household debt.

Figure 2.5: Baseline model dynamic multipliers of a shock to credit growth of 1 percentage point

The figures show that an exogenous shock in one of the markets is propagated and amplified through an endogenous feedback mechanism. Figure 2.5 shows that a positive exogenous shock in the credit growth by one percentage point will increase housing price

FIGURE 2.6: Baseline model dynamic multipliers of a shock to housing price growth of 1 percentage point

growth by 0.86 percentage points at the time of the shock, which equals the instantaneous impact on housing price growth reported in Table 2.5. The increase in housing prices leads to a further increase in credit growth in the subsequent period, as the collateral value has increased. This again induces further growth in housing prices and credit in a process that continues for about two years before the equilibrium correction term dominates and the effect of the shock gradually dissipates. In the long-run, there is of course no change in neither of the growth rates, but we see that the levels of both variables have stabilized at a higher level in accordance with the finding of a long-run interaction between housing prices and credit in Section 2.5. Shocking housing price growth (see Figure 2.6) yields qualitative effects that are similar to the above described effects, and will of course not change any of the growth rates in the long-run.

A shock to one of the exogenous variables in the system will have similar effects as is shown in Figure 2.7. A one percent increase in disposable income will lead to a growth in both housing prices and credit, which is reinforced by the feedback between the two variables. The dynamic process clearly indicates that the relationship between housing prices and credit is mutually self-reinforcing. First, a higher income leads to increased property valuations, which raises the value of the collateral. This spills over to the credit market, stimulating housing prices further, and so on. As the cumulative multipliers illustrate, both housing prices and credit continue to grow before the growth rates eventually return to zero. This has of course lead to a new equilibrium price level

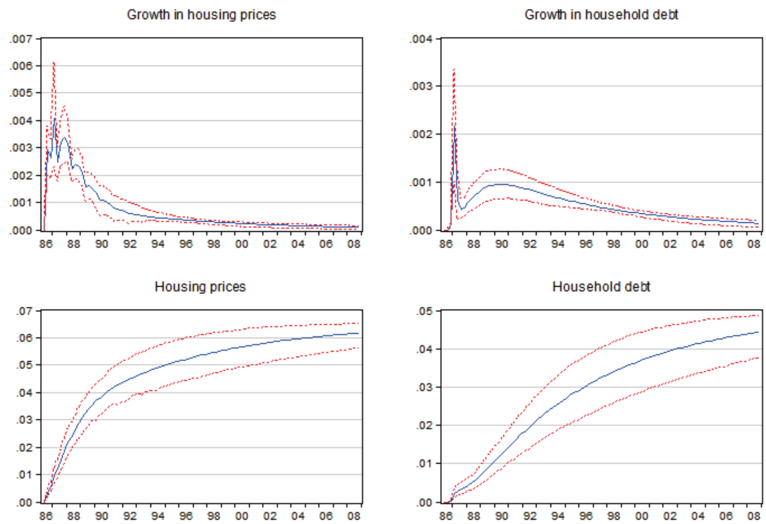

FIGURE 2.7: Baseline model dynamic multipliers of an increase in real disposable household income by 1 percent

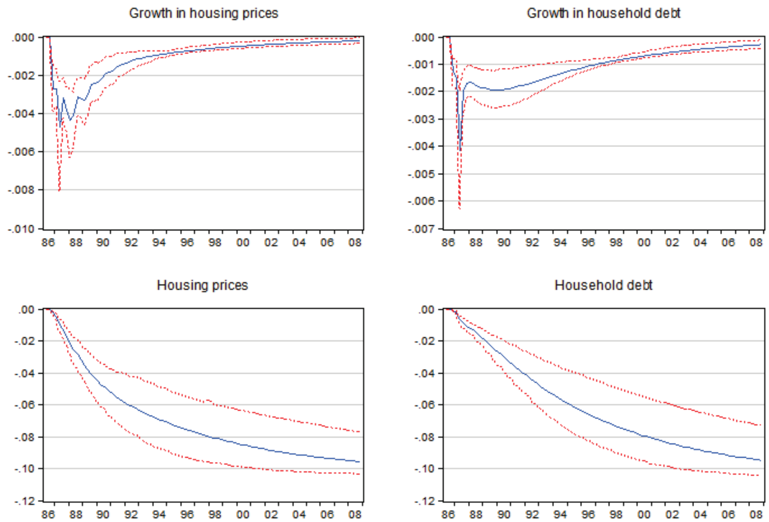

FIGURE 2.8: Baseline model dynamic multipliers of a shock to the interest rate of 1 percentage point

and a higher fundamental value for the credit variable, as seen from the lower part of the figure. An increase in disposable income, which is one of the long-run determinants of housing prices, will change housing prices and credit period after period until they have adjusted to their new long-run equilibrium level.

Figure 2.8 shows the simulated responses to a one percentage point increase in the real interest rate. This reduces both housing prices and credit growth in the short-run. In the long-run, both housing prices and household debt converge to new and lower equilibrium levels (lower part of the figure), which shows that the model implies interest rates effects on housing prices even though the interest rate does not enter the short nor the long-run equations for housing prices directly.

2.7.2 Dynamic multipliers: An extended model

In this section we augment the core model above with a small model for the supply side of the housing market. These equations are lifted out of the macroeconometric forecasting model KVARTS, which is an operative and relevant model for the Norwegian economy. The supply side model captures the feedback from housing prices to the investments in new houses, which again affects the housing stock and therefore is expected to dampen the dynamic effects found in the previous subsection. The housing supply model is reestimated on our sample and a brief description of the supply side model, along with the estimated coefficients, are given in Appendix 2.A. Figure 2.9 and Figure 2.10 illustrate the dynamic impact of a one percentage point increase in credit growth and housing price growth when the supply side is taken into account.

Though the short-run effects are very similar to those for the baseline model, we see that the effects of the shocks on the growth rates die off more quickly when taking into account that the investment activity responds to changes in housing prices. While in the baseline model a 1 percentage point increase in credit growth still has a great effect on the housing price growth after 4 years, we find that the estimated effect on housing price growth is zero in the extended model after the same period. It follows that also the long-run impact on the levels of housing prices and credit are much reduced, as is seen from the graphs in the middle part of Figures 2.9 and 2.10. In the long-run, we see the expected convergence to a new equilibrium with higher housing prices and a greater housing stock.

In Figure 2.11, we have graphed the simulated responses when we increase household disposable income by 1 percent. Again, it is clear that including the supply side dampens the effects relative to those reported in the previous section. In the baseline model, this income shock leads to an increase in housing prices of more than 4 percent after 4 years, and in the long-run the estimated effect on housing prices is around 6 percent. This contrasts the extended model, where the effect on housing prices after 4 years is around 3 percent. At this point, the effect gradually declines, as the investment activity increases. In the long-run, we find that housing prices have increased by 0.5 percent, which is half of the initial increase in income. Household debt is found to increase by 1 percent, meaning that the long-run effect on debt will equal the initial shock to income.

The final figure (Figure 2.12) shows the effect of an increase in the real interest rate of one percentage point. Again, the short-run responses are similar to those in the baseline model, while the long-run effects are much reduced. It should be noted that the disposable income variable includes net interest rate income, which is negative on aggregate for the households. Thus, if we had used a larger model, where also disposable income had been modelled, the simulated interest rate effect would be stronger.

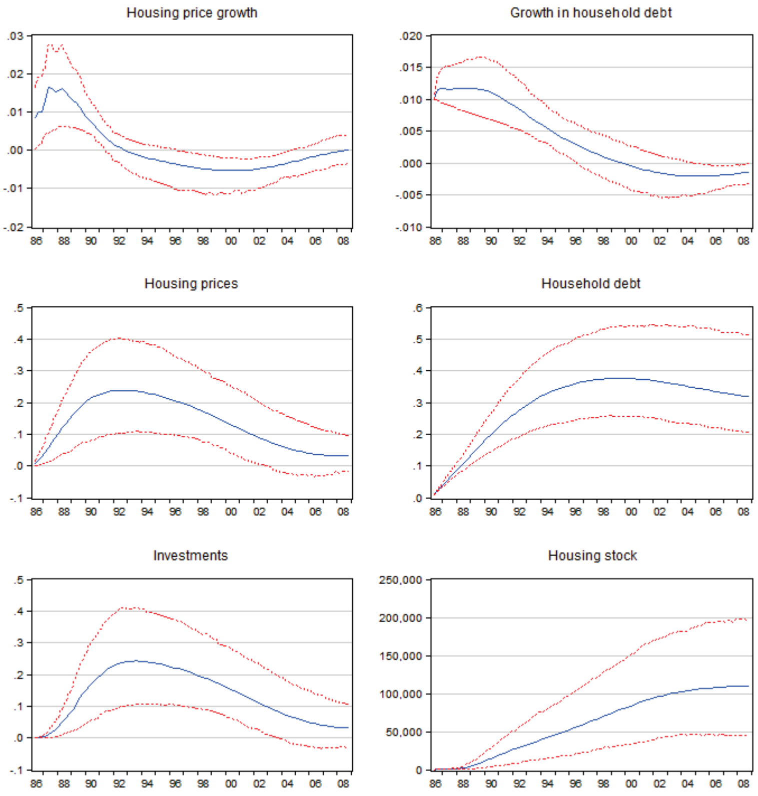

FIGURE 2.9: Dynamic multipliers of a shock to credit growth of 1 percentage point in the extended model.

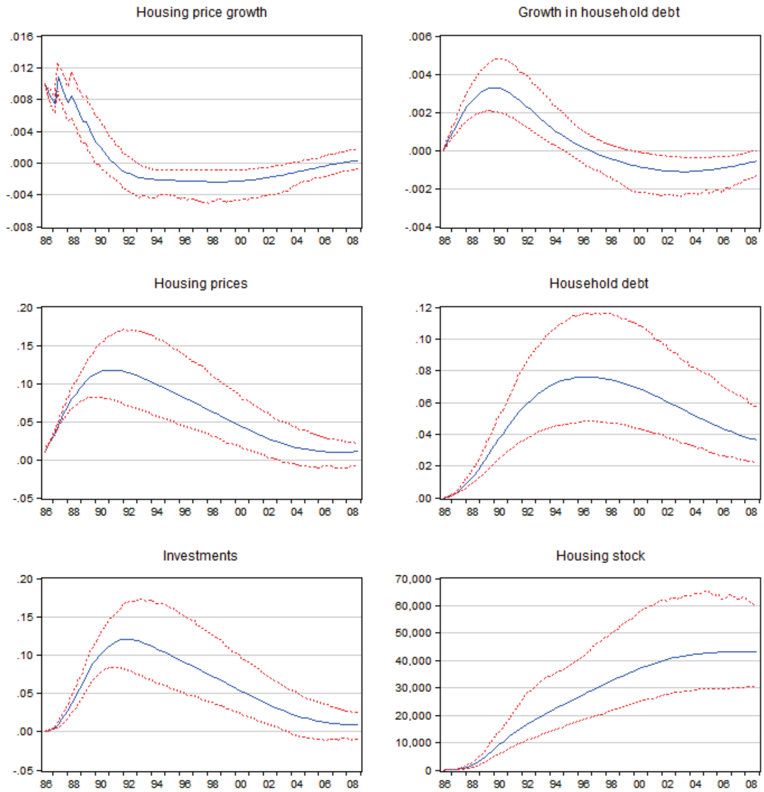

FIGURE 2.10: Dynamic multipliers of a shock to housing price growth of 1 percentage point in the extended model.

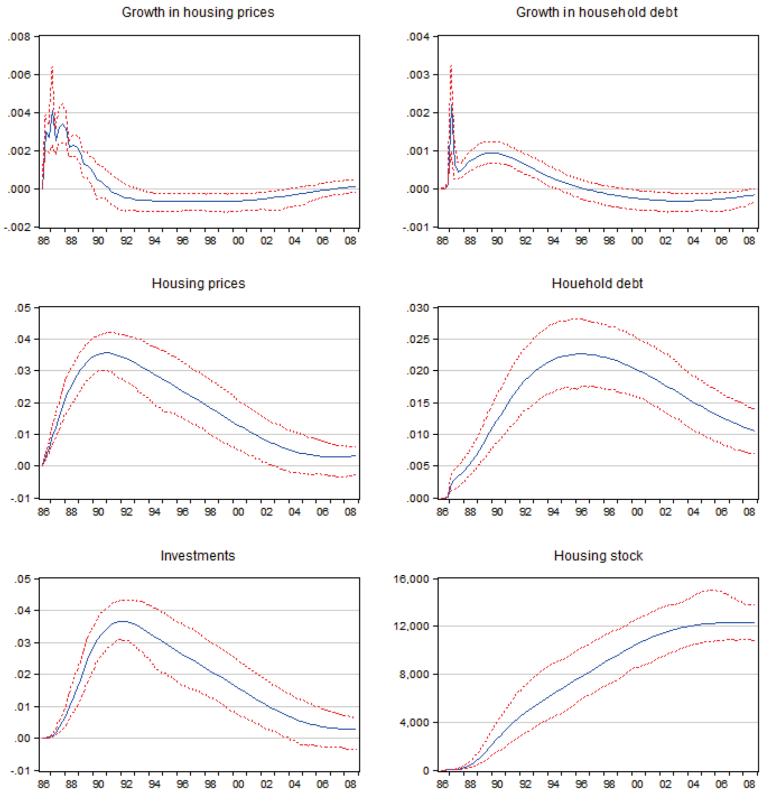

Figure 2.11: Dynamic multipliers of an increase in real disposable household income of 1 percent in the extended model.

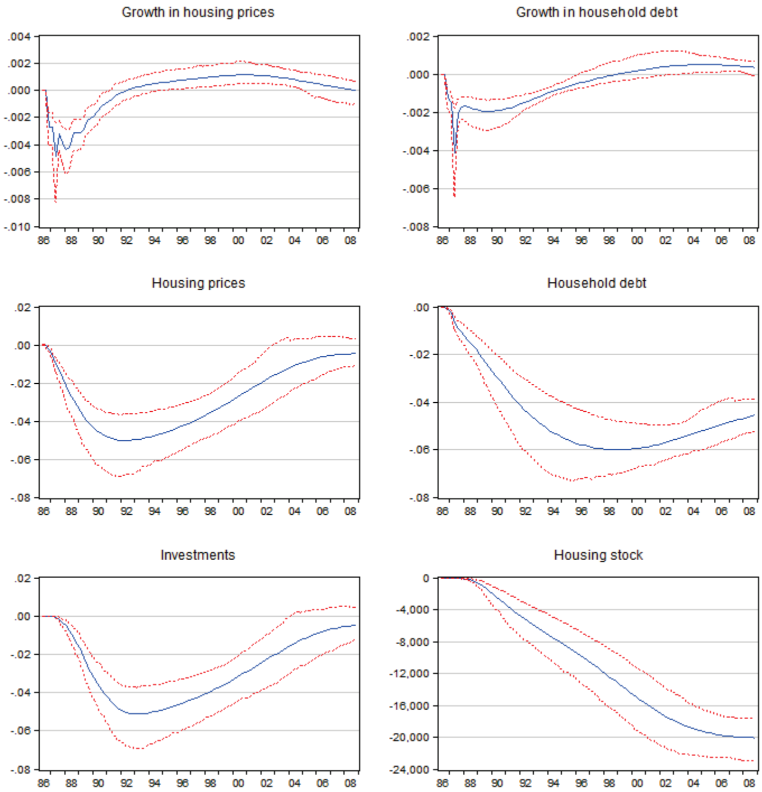

FIGURE 2.12: Dynamic multipliers of an increase in the real interest rate of 1 percentage point in the extended model

2.8 Robustness: Estimating the model on an extended sample

With the benefit of having access to a four more years of data than we had when we first started working on the Anundsen and Jansen (2013b) paper, we have reestimated the short-run dynamics of the model for every quarter between the period 2008q4–2012q4. In addition to having an extended data set, there have also been revisions to the data we originally used. Thus, such a reevaluation of the model is useful to explore the robustness of our results. The recursive coefficient estimates are reported in Figure 2.13 and 2.14.

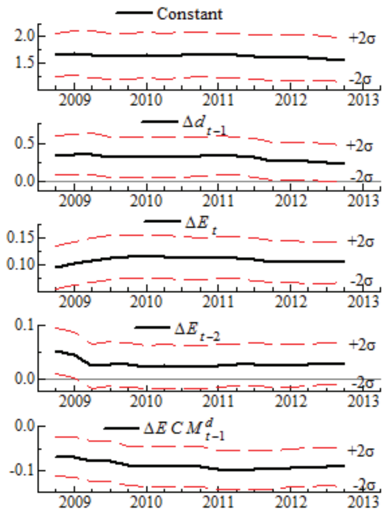

FIGURE 2.13: Recursively estimated coefficients for Δph equation from the "baseline" model, 2008q4–2012q4

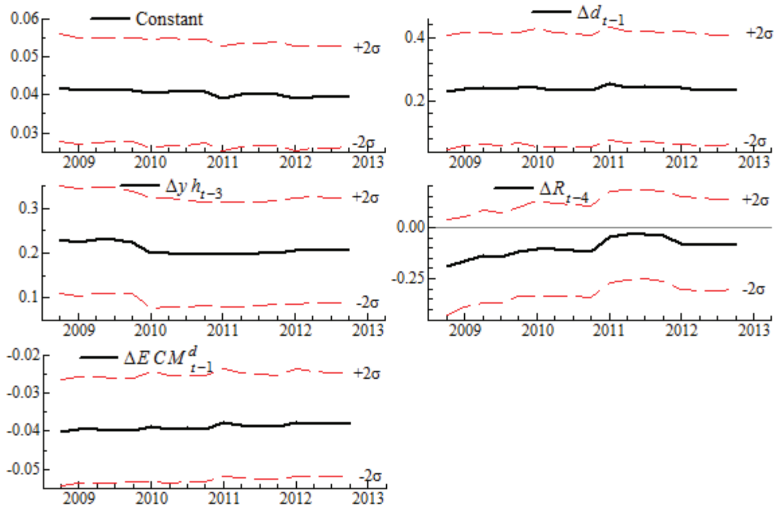

FIGURE 2.14: Recursively estimated coefficients for Δd equation from the "baseline" model, 2008q4–2012q4

It is clear that all the coefficients are stable when estimated recursively. The same recursive estimates for the model *without* short-run price homogeneity are reported in Figure 2.E.1 and 2.E.2 in Appendix 2.E and they show the same picture. This is a reassuring finding, and is particularly important if the model is to be used for forecasting purposes. Having a good forecasting model for housing prices and credit seems imperative both in order to monitor the development in the financial system and to increase the forecasting accuracy of key macroeconomic variables such as consumption and investments. In fact, preliminary results (Anundsen and Jansen, 2013a) show that the *ex ante* forecasts from the model documented in this chapter fares well against alternative forecasting models, such as autoregressive, vector autoregressive and random walk models.

Finally, and again with the benefit of having access to more data, Figure 2.15 and 2.16 show the forecasts for the model *with* and the model *without* short-run price homogeneity for the period 2008q4–2012q4, i.e. adding five more observations to the forecasting horizon relative to what we did in Section 2.6

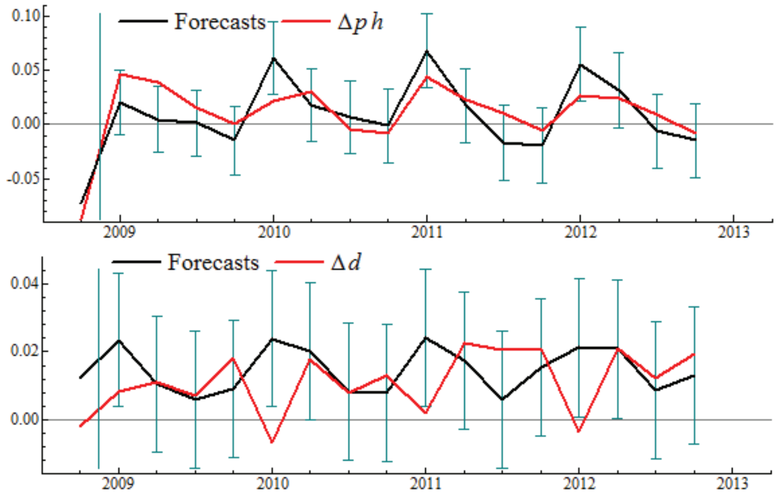

FIGURE 2.15: Ex ante forecasts from the "baseline" model, 2009q1–2012q4

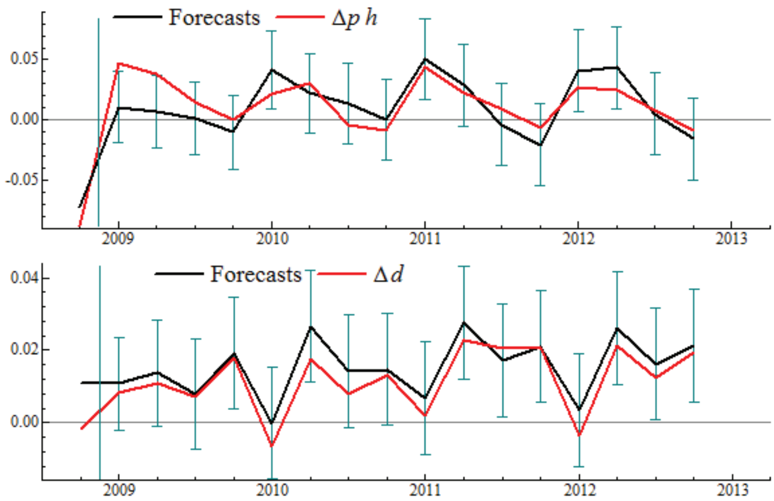

FIGURE 2.16: Ex ante forecasts from the model *without* short-run price homogeneity, 2009q1–2012q4

Also for the last year, we see that the credit forecasts produced by the baseline model are outside their confidence bounds in the first quarter. However, the model where we have derestricted the short-run price homogeneity does a far better job, which lends support to our argument that including short-run inflation effects in the credit equation may be important for forecasting purposes. In conclusion, it seems that the model passes the stability tests when evaluated on an extended sample.

2.9 Conclusion

Using cointegration analysis, this study documents the importance of jointly estimating long-run interactions between house prices and household debt. Furthermore, estimating these variables in a vector error-correction system also yields better estimates of short-run interactions and dynamic responses. We find evidence that household income is weakly exogenous with respect to other long-run housing-related variables. Along with other tested constraints on coefficients, we use this finding to estimate a more parsimonious system of household debt and house prices.

In particular, we find that house prices depend on household borrowing, real disposable income and the housing stock in the long-run, whereas real household debt is driven by the value of housing capital (housing prices times the housing stock), the real interest rate and the housing turnover. Housing prices and household debt are mutually dependent as both appear in the long-run equation for the other. This suggests that there are feedback effects between the two in the long-run. That said, housing prices are equilibrium correcting to deviations from both long-run equations, whereas household debt adjusts only to disequilibria in the credit market.

Second, we embed the long-run equations from the cointegration analysis in a simultaneous system explaining the changes in housing prices and debt, following a general-to-specific strategy. The equations are estimated simultaneously by full information maximum likelihood methods and insignificant variables are removed stepwise from the two equations. The estimation results suggest that the credit aggregate is important for housing price dynamics, but that housing prices only affect household borrowing through the equilibrium correction term.

Third, a consumer confidence indicator measuring households' expectations concerning future developments in their own economy as well as the Norwegian macro economy are incorporated into our framework. This variable explicitly picks up expectations about future economic conditions and is shown to enter significantly in the housing price equation in the short-run.

Finally, the analysis of the dynamic multipliers provides clear evidence for the existence of a credit-housing price spiral in Norway. Higher housing prices result in higher credit growth due to collateral effects, which again spurs housing price growth and so on, showing that there indeed is a financial accelerator at work. Incorporating a model of the supply side of the housing market dampens the dynamic responses of housing prices and credit to all shocks considered here. This highlights the importance of accounting for construction, as well as credit factors, in modeling housing cycles.

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Appendix 2.A The supply side

The equations describing the supply side of the housing market in Section 2.7 are lifted out of the Statistics Norway quarterly forecasting model, KVARTS (Eika and Moum (2005)) and reestimated on the current sample (1986q2-2008q4). In KVARTS, the supply of housing is modelled by considering housing starts measured in square meters. Housing starts serve as a leading indicator for the development in housing investments, which eventually become new houses and add to the housing stock.

In a long-run perspective, new housing starts are modeled according to the q-theory of investments, where a one percent increase in either housing prices or a one percent decrease in construction costs lead to a one percent increase in housing starts. This implies that a proportional increase in construction costs and housing prices will have no long-run effect on the supply of new houses. Letting S denote housing starts, PJ denote real construction costs and PH denote real housing prices, the reestimated equation for housing starts is given by (absolute t-values reported under the point estimates).

$$\begin{aligned} \Delta logS_t &= \underset{(4.90)}{0.41} \Delta logS_{t-4} - \underset{(4.28)}{0.26} \left(logS_{t-1} - logPH_{t-1} + logPJ_{t-1}\right) \\ &\quad + dummies \\ R^2 &= 0.77 \end{aligned} \tag{2.A.1}$$

In addition to the equilibrium correction term, the model contains an autoregressive part as well as an impulse dummy for the second quarter of 2002 and a set of seasonal dummies for the first three quarters. The re-estimated coefficients are almost unchanged from the version used in KVARTS, which is reassuring.

Since it takes time for a newly started building project to get finished, it is assumed that a change in housing starts will lead to a flow of investments for several years. In KVARTS this adjustment is assumed to take 12 quarters and the relationship linking investments and housing starts is given by the following equation:

$$\Delta logIH = \Delta logJ + seasonals \tag{2.A.2}$$

where IH denotes housing investments, which grow proportionally with a weighted average of housing starts over the last 12 quarters, J. Also the coefficients for the seasonal dummy variables in equation (2.A.2) are reestimated when we construct the model used for simulations in Section 2.7. The weighted average of housing starts is given by the following identity .

$$J_t = 0.3124 * S_t + 0.2455 * S_{t-1} + 0.1672 * S_{t-2} + 0.1125 * S_{t-3} + 0.0702 * S_{t-4}$$
$$+ 0.0407 * S_{t-5} + 0.0235 * S_{t-6} + 0.0131 * S_{t-7} + 0.0074 * S_{t-8} + 0.0043 * S_{t-9}$$
$$+ 0.0021 * S_{t-10} + 0.009 * S_{t-11} + 0.002 * S_{t-12}$$

Finally, the housing stock is determined by a law of motion of capital accumulation:

$$H_t = (1 - \delta)H_{t-1} + IH_t$$

where δ is the rate of depreciation of the housing stock. As is evident from this brief presentation of the supply side, the model used for simulation in Section 2.7 captures spill-overs from housing prices to the construction sector, which, as shown in the simulation exercises, dampens the long-run effect of shocks on housing prices and credit.

Appendix 2.B Data definitions

All data are seasonally unadjusted and measured on a quarterly basis. Except for the interest rate and the consumer confidence indicator all variables are transformed to log scale in the empirical analysis. Variable definitions and a brief description of the data are listed below.

pc: The consumption deflator in the National Accounts. Source: Statistics Norway.
ph: Hedonic housing price index measuring average housing prices in Norway. The index is calculated on the basis of data on sales in the second hand market. Statistics Norway officially started publishing housing price data in 1992. Prior to 1992 an unofficial index based on similar sources and compiled at Statistics Norway is used. The housing price index is deflated by pc. Source: Statistics Norway.
d: Total amount of outstanding gross household debt. Deflated by pc. Source: Statistics Norway.
yh: Households' disposable income, excluding equity income. Deflated by pc. Source: Statistics Norway.
h: Real housing stock measured in fixed prices. Measures the total stock of housing in Norway and is calculated according to the perpetual inventory method. Source: Statistics Norway.
th: The housing turnover measures the number of housing transactions. Source: Statistics Norway.
E: The expectations variable is taken from TNS Gallup and can be seen as a consumer confidence indicator. It is based on a survey, where average score can range between -100 and 100. In this paper, we have normalized the variable to lie between -1 and 1. The indicator measures households expectations concerning the state of the economy and the development in their personal economy. Source: TNS-Gallup.
i: Nominal interest rate paid by households on loans in private financial institutions. Source: Statistics Norway.
p: Consumer Price Index. Source: Statistics Norway.
π*:* Annual inflation rate ($\Delta_4 p$).
τ*:* Capital tax rate. After a tax reform in 1992 τ has been constant at 0.28. Source: Statistics Norway.
R: Real after-tax interest rate ($i * (1 - \tau) - \pi$).

Variables used in Appendix 2.A:
S: Housing starts (square meters). Source: Statistics Norway.
J: Weighted sum of housing starts (square meters).
IH: Investments in housing, measured at fixed prices. Source: Statistics Norway.

PJ: Price index for construction costs, deflated by *pc*. Source: Statistics Norway.

δ*:* rate of depreciation of the housing stock.

Appendix 2.C Tables

TABLE 2.C.1: Tests for the order of integration[a]

	ADF			PP	KPSS		
Testing levels							
Variable	$t-ADF$	5%	$Adj.t-stat$	5%	LM	5%	Characteristics[b]
ph	−2.37	−3.46	−1.32	−3.46	0.27	0.146	t
d	−3.77	−3.46	−0.69	−3.46	0.27	0.146	t
h	−2.76	−3.46	−0.78	−3.46	0.22	0.146	t
yh	−0.98	−3.46	−5.18	−3.46	0.31	0.146	t
th[c]	−3.21	−3.46	−7.74	−3.46	0.14	0.146	t
r	−3.58	−3.46	−3.5	−3.46	0.13	0.146	t
E[d]	−1.80	−3.46	−2.15	−3.46	0.08	0.146	t
Testing first differences							
Δph	−2.07	−2.89	−5.99	−2.89	0.25	0.46	i
Δd	−1.77	−2.89	−5.35	−2.89	0.3	0.46	i
Δh	−2.20	−2.89	−1.84	−2.89	0.29	0.46	i
Δyh	−4.25	−2.89	−27.05	−2.89	0.44	0.46	i
Δth	−8.71	−2.89	−21.91	−2.89	0.11	0.46	i
Δr	−11.11	−2.89	−10.73	−2.89	.10	0.46	i
ΔE	−5.12	−2.89	−7.55	−2.89	0.28	0.46	i
Testing second differences							
$\Delta^2 ph$	−4.62	−2.89	–	–	–	–	i
$\Delta^2 d$	−13.28	−2.89	–	–	–	–	i
$\Delta^2 h$	−2.548	−2.89	−11.41	−2.89	–	–	i

[a] While the PP and KPSS tests are performed in EViews, we run the ADF test in PcGive since this allow us to include seasonal dummies in the test regression. The variables for which we have included seasonal dummies in the test regressions are housing prices, disposable income and the turnover, as they all display a clear seasonal pattern. When inspecting this table, it is important to keep in mind that while the ADF test and the PP test have non-stationarity as the null, the KPSS test has stationarity as the null.

[b] The different characteristics are: Including both trend and intercept (t) or only an intercept (i) in the test regression.

[c] The turnover is only collected from 1985q1, which means that with 8 lags in the ADF regression, the sample starts in 1987q2.

[d] For the expectations variable we only have data for the period from 1992q3 and the variable is set to 0 in the period prior to this in the empirical analysis. For the tests for the order of integration, we use the period for which we have observations.

Table 2.C.2: Lag reduction for the exogenous variables in the unrestricted VAR [a,]

Lags	log likelihood	SC	HQ	AIC
5	869.13433	-14.194	-15.824	-16.926
4	866.47195	-14.433	-15.964	-16.999
3	860.07987	-14.590	-16.022	-16.991
2	857.56754	-14.832	-16.166	-17.067
1	854.16023	-15.055	-16.290	**-17.124**
0	845.28489	**-15.157**	**-16.293**	-17.061
Tests of lag reduction				
5 to 4	F(6,112) =	0.55420 [0.7658]		
5 to 3	F(12,148) =	0.96638 [0.4836]		
5 to 2	F(18,158) =	0.83006 [0.6629]		
5 to 1	F(24,163) =	0.81618 [0.7127]		
5 to 0	F(30,165) =	1.0756 [0.3722]		
4 to 3	F(6,116) =	1.4069 [0.2178]		
4 to 2	F(12,153) =	0.98362 [0.4670]		
4 to 1	F(18,164) =	0.91767 [0.5582]		
4 to 0	F(24,168) =	1.2251 [0.2269]		
3 to 2	F(6,120) =	0.55985 [0.7615]		
3 to 1	F(12,159) =	0.66799 [0.7801]		
3 to 0	F(18,170) =	1.1519 [0.3071]		
2 to 1	F(6,124) =	0.78849 [0.5806]		
2 to 0	F(12,164) =	1.4710 [0.1398]		
1 to 0	F(6,128) =	2.1855[0.0485]*		
Estimation period:	1986q2-2008q4			

[a] Endogenous variables: Real housing prices, real household debt and real disposable income. Restricted variables: Real interest rate after tax, housing turnover, housing stock and a linear trend. Unrestricted variables: Constant and seasonal dummies.

Table 2.C.3: Loading factors for the models reported in Table 2.4

Loading	Panel 1	Panel 2	Panel 3	Panel 4	Panel 5
$\alpha_{1,ph}$	-0.82 0.21	-0.76 0.20	-0.21 0.04	-0.22 0.04	-0.24 0.04
$\alpha_{1,d}$	-0.44 0.14	-0.40 0.13	-0.06 0.02	-0.08 0.03	-0.1 0.03
$\alpha_{2,ph}$	-0.13 0.11	-0.13 0.11	0 $--$	0 $--$	0 $--$
$\alpha_{2,d}$	-0.11 0.07	-0.10 0.07	-0.04 0.01	-0.05 0.01	-0.04 0.01
$\alpha_{3,ph}$	0.42 0.15	0.40 0.14	-0.04 0.03	-0.05 0.03	0 $--$
$\alpha_{3,d}$	0.31 0.1	0.29 0.09	0.002 0.02	-0.01 0.02	0 $--$

Note: This table reports the estimated loading factors (equilibrium correction coefficients) obtained when we impose the various overidentifying restrictions on our two cointegrating vectors, confer Table 2.4 for the estimated cointegrating vectors. Below the point estimates are the estimated standard errors.

Table 2.C.4: Augmented Dickey-Fueller tests for structural residuals[a]

Variable	t-ADF	5%-critical value	lags	trend	seasonal dummies
$\varepsilon_{\Delta ph}$	-8.846	-2.89	0	No	No
$\varepsilon_{\Delta d}$	-7.945	-2.89	1	No	No

[a] The residuals from the short run system (confer Table 2.5) are tested over the period 1988q3-2008q4 since we only obtain data for the error correction terms from 1986q2.

Appendix 2.D Equation-by-equation modeling

Adopting a single equation approach one would take the system represented by equation (2.9) and (2.10) as a starting point. This approach precludes any formal treatment of identification, but may possibly give reasonable results if the simultaneity bias is not large. We have used the automated multipath search algorithm Autometrics (see Doornik (2009) and Doornik and Hendry (2009b)) to reduce the dimensionality of each equation. An obvious advantage with this algorithm is that it is very little path dependent as it does a multipath search. However, the benefit from this might be outweighed by the fact that it does not allow us to take care of the simultaneity from the onset by doing a full fledged system analysis at each step in the reduction process. The results from this single equation general to specific approach are documented in Table 2.D.1 and Table 2.D.2 for the housing price and credit equation, respectively.

Table 2.D.1: Short run dynamics obtained by Autometrics for housing price equation[a]

Variable	Coefficient	t-value
$Constant$	1.23	6.78
Δd	0.61	3.85
Δph_{t-4}	0.41	4.93
Δt_{t-3}	0.05	2.55
Δr_{t-4}	-0.38	2.06
ΔE_t	0.095	4.54
ΔE_{t-1}	0.096	4.40
ΔE_{t-2}	0.05	2.17
ecm^{ph}_{t-1}	-0.07	3.81
ecm^{d}_{t-1}	-0.14	6.80
$CSeasonal_t$	-0.006	0.496
$CSeasonal_{t-1}$	-0.007	0.65
$CSeasonal_{t-2}$	-0.009	0.999
σ	0.0141	
R^2	0.82	
$Adj.R^2$	0.80	
Diagnostics[b]	Test statistic	Value [p-value]
AR 1-5 test:	$F(5,73) =$	0.4789 [0.7909]
ARCH 1-4 test:	$F(4,83) =$	0.4462 [0.7749]
Normality test:	$\chi^2(2) =$	1.5603 [0.4583]
Hetero test:	$F(21,69) =$	1.3658 [0.1672]
Estimation Method	OLS (Autometrics with p-value = 0.05)	
Sample	1986q2-2008q4	

[a] Absolute t-values are reported.
[b] See Doornik and Hendry (2009b).

Table 2.D.2: Short run dynamics obtained from Autometrics for the credit equation[a]

Variable	Coefficient	t-value
$Constant$	-0.73	10.6
Δph_t	0.30	7.06
Δph_{t-4}	-0.12	2.64
Δy_{t-2}	-0.15	3.10
ΔE_{t-1}	-0.04	2.45
Δr_{t-3}	-0.24	2.34
ecm^{ph}_{t-1}	0.09	10.8
$CSeasonal_t$	-0.004	1.16
$CSeasonal_{t-1}$	-0.004	1.50
$CSeasonal_{t-2}$	-0.01	4.07
σ	0.009	
R^2	0.72	
$Adj.R^2$	0.69	
Diagnostics[b]	Test statistic	Value [p-value]
AR 1-5 test:	$F(5,76) =$	1.4959 [0.2011]
ARCH 1-4 test:	$F(4,83) =$	0.7501 [0.5608]
Normality test:	$\chi^2(2) =$	4.9864 [0.0826]
Hetero test:	$F(15,75) =$	0.8092 [0.6641]
Estimation Method	OLS (Autometrics with p-value = 0.05)	
Sample	1986q2-2008q4	

[a] Absolute t-values are reported.
[b] See Doornik and Hendry (2009b).

The results in Table 2.D.1 and Table 2.D.2 reveal some differences as compared to our preferred model. We note that both variables enter contemporaneously in both equations. Also, we observe that the income variable and the expectations variable are both highly significant in the credit equation with negative signs, which is not plausible *a priori* . Let us now turn to the two equations when they are estimated simultaneously to take care of potential endogeneity problems. Results are displayed in Table 2.D.3.

TABLE 2.D.3: System estimation of the specifications obtained by Autometrics (equation-by-equation)[a]

	Real housing prices		Real household debt	
Variable	Coefficient	t-value	Coefficient	t-value
Constant	1.00	3.78	−0.73	10.5
Δd_t	−0.26	0.49	–	–
Δph_t	–	–	0.32	5.50
Δph_{t-4}	0.36	3.65	−0.13	2.57
Δyh_{t-2}	–	–	−0.15	3.05
ΔE_t	0.12	3.88	–	–
ΔE_{t-1}	0.10	3.95	−0.04	2.48
ΔE_{t-2}	0.05	1.75	–	–
Δr_{t-3}	–	–	−0.24	2.37
Δr_{t-4}	−0.51	2.36	–	
Δt_{t-3}	0.06	2.50	–	
ECM^{ph}_{t-1}	−0.11	3.34	0.09	10.6
ECM^{d}_{t-1}	−0.10	3.85	–	–
Dummy, q1	−0.01	0.75	−0.005	1.26
Dummy, q2	−0.009	0.73	−0.004	1.55
Dummy, q3	−0.02	1.61	−0.01	4.07
Sargan	$\chi^2(43) =$	40.323 [0.5881]		
Log likelihood	567.99			
σ	0.016		0.0086	
Diagnostics[b]	Test statistic	Value [p-value]		
Vector SEM-AR 1-5 test:	$F(20, 138) =$	0.7944[0.7168]		
Vector Normality test:	$\chi^2(4) =$	4.7544[0.3134]		
Vector Hetero test:	$F(183, 81) =$	1.0260[0.4557]		
Estimation Method	FIML			
Sample	1986q2-2008q4			

[a] Absolute t-values are reported.
[b] See Doornik and Hendry (2009a).

The credit equation remains almost unaltered, while the housing price equation changes dramatically. First of all, the credit variable which is positive and highly significant in the single equation model has now changed sign and is insignificant. Also, the loadings have changed. As a final check of this model, we will explore how the implied dynamics of the system to a permanent increase in real disposable income would be. We follow exactly the same set up as in Section 2.7.1 of the paper and the dynamic multipliers are graphed in Figure 2.D.1.

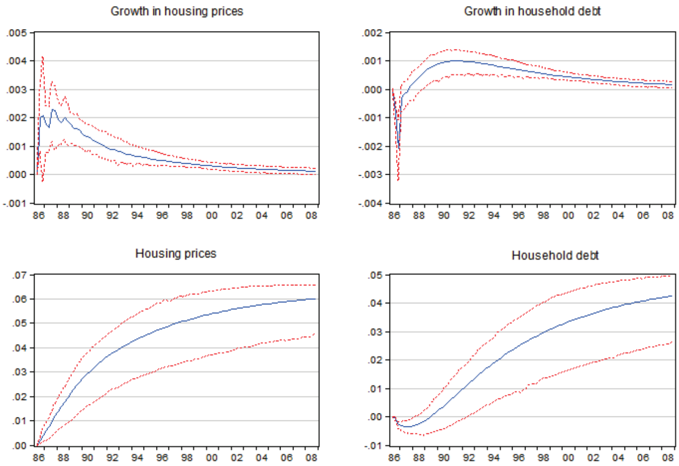

FIGURE 2.D.1: The alternative model: Dynamic multipliers of a 1 percent increase in real disposable household income.

Based on the dynamic multipliers from this alternative model, we see that it implies a negative response to household borrowing of an increase in income in the short run, which seems unreasonable from an economic theory point of view. Also, the credit effect on housing prices changes sign and turns out insignificant, though it was positive and highly significant in the single equation case. Furthermore, we observe relatively big changes in the loadings in the housing price equation. On this background we conclude that this model is inferior to the one from the simultaneous model design reported in Table 2.5 in Section 2.6.2 of the paper.

Appendix 2.E Model without short-run price homoegeneity

With reference to the forecasting exercise in Section 2.8 of the paper, this section discusses a version of the model, where we de-restrict the assumption of short-run price homogeneity. To see whether the forecast failures for the credit growth in 2010q1 and 2011q1 (confer Figure 2.3 in the paper) may be due to the extremely cold winters, which lead to an extraordinary jump in electricity prices in each of the two quarters, we re-estimated the model for the case where short-run price homogeneity is relaxed. As shown in the paper (see Figure 2.4), this improves the forecasting accuracy of the model – and in particular the credit forecasts. The estimation results underlying those forecasts are reported in Table 2.E.1.

TABLE 2.E.1: Short-run dynamics from the model *without* short-run price homogeneity[a]

	Real housing prices		Real household debt	
Variable	Coefficient	t-value	Coefficient	t-value
Constant	1.617	7.90	0.023	4.83
Δd_t	0.696	3.78	-	-
Δd_{t-1}	-	-	0.560	7.68
Δd_{t-3}	0.355	2.69	-	-
Δph_{t-4}	0.394	5.07	-	-
Δyh_{t-3}	-	-	0.084	1.99
ΔE_t	0.102	5.12	-	-
ΔE_{t-1}	0.100	4.76	-	-
ΔE_{t-2}	0.045	2.05	-	-
ΔR_{t-4}	-	-	-0.088	1.13
Δpc_t	-	-	-0.720	9.25
Δpc_{t-1}	-	-	0.528	5.89
ECM^{ph}_{t-1}	-0.172	7.86	-	-
ECM^{d}_{t-1}	-0.071	4.26	-0.025	4.63
Dummy, q1	0.025	3.87	-0.016	4.37
Dummy, q2	0.024	4.27	0.007	2.52
Dummy, q3	0.013	2.31	-0.019	7.36
Sargan		$\chi^2(48) =$	44.68 [0.6099]	
Log likelihood		603.68		
σ	0.0137		0.0064	
Diagnostics[b]		Test statistic	Value [p-value]	
Vector EGE-AR 1-5 test:		F(20,138)	0.50 [0.96]	
Vector Normality test:		$\chi^2(4)$	36.17 [0.00]	
Vector hetero test:		F(195,69)	0.67 [0.98]	
Estimation Method	FIML			
Sample	1986q2-2008q4 ($T = 91$)			

[a] Absolute t-values are reported.
[b] See Doornik and Hendry (2009a).

We started by including the current and first lag of the change in the price deflator (Δpc) in both equations. However, these variables were only significant in the credit equation, and were therefore excluded from the housing price equation. As seen, the inclusion of Δpc_t and Δpc_{t-1} in the credit equation only has minor effects on the estimated parameters of the housing price equation, while the estimates of the credit equation are somewhat changed. That said, it seems to be changed for the better, since – as is evident from inspecting the table – derestricting short-run price homogeneity improves the fit of the credit equation. Furthermore, both the current and lagged value are highly significant, and come with opposite signs. In fact, we can not reject the hypothesis that the two coefficients are equal in absolute value, i.e. suggesting that these terms are measuring a surprise inflation ($\Delta^2 pc_t = \Delta pc_t - \Delta pc_{t-1}$). This gives additional credence to our conjecture that the forecast failures in 2010q1 and 2011q1 are due to an unexpected increase in electricity prices.

Recursive estimates for this model for the period 2008q4–2012q4 are displayed in Figure 2.E.1 and 2.E.2.

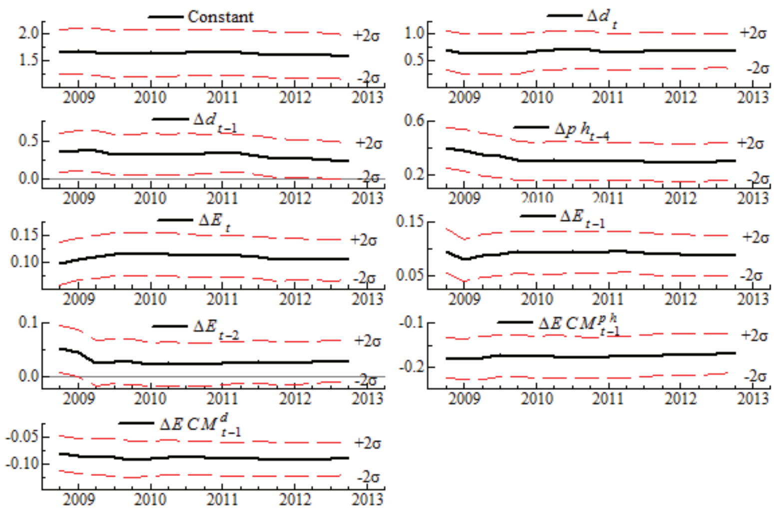

FIGURE 2.E.1: Recursively estimated coefficients for Δph equation from the model *without* short-run price homogeneity, 2008q4–2012q4

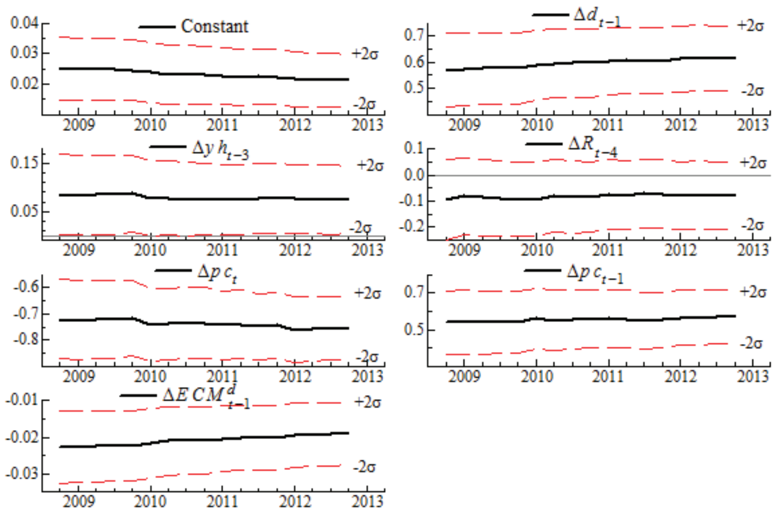

FIGURE 2.E.2: Recursively estimated coefficients for Δd equation from the model *without* short-run price homogeneity, 2008q4–2012q4